ESSAYS ON HOUSEHOLD CONSUMPTION AND FINANCE

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Declaration

I hereby declare that this thesis is my original work and it has been written by me in its entirety. I have duly acknowledged all the sources of information which have been used in the thesis.

This thesis has also not been submitted for any degree in any university previously.

Lai Xiongchuan
26 May 2015

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Summary

This thesis consists of three essays that have the common theme of examining the connection between household consumption/finance choices and macroeconomic performance. The first essay examines how housing consumption and elasticity of housing supply could affect equity risk premium and housing risk premium. It provides both theoretical predictions and empirical evidence for the connection between asset risk premia and elasticity of housing supply. The second essay investigates how the cross-sectional variations of housing supply elasticity have implications on households' portfolio composition. It finds that households living in areas with less elastic housing supply invest more in stocks for the purpose of hedging housing consumption risk. Lastly, the third essay examines competitive consumption and labor supply behavior of young males in China in connection to sex ratio imbalance.

The first essay extends the housing consumption-based asset pricing model in Piazzesi *et al.* (2007) to a production economy, where housing consumption is endogenous with respect to both aggregate productivity shocks and housing supply elasticity. The role of housing as a consumption good in shaping asset risk premia is re-examined. In contrast to the exchange-economy case where the presence of housing introduces an independent consumption composition risk to elevate asset risk premia, adding housing to the consumption basket in a production economy introduces a substitution benefit that mitigates consumption

risk and lowers asset risk premia. Moreover, lower housing supply elasticity makes housing price more volatile in response to productivity shocks, thus reducing the equity risk premium via enhanced substitution benefit but increasing the housing risk premium via elevated consumption risk. Empirical analysis using land share of home value as proxy for aggregate housing supply inelasticity in the economy shows that a lower housing supply elasticity predicts lower excess stock returns but higher excess housing returns, especially in the long-horizon (6-12 quarters) return forecasts. Besides clarifying the role of housing in consumption-based asset pricing models, these findings also provide an alternative explanation for the declining equity risk premium observed in recent decades.

The second essay uses geographic variation of the housing supply elasticity to account for housing consumption risk and investigates the influence of such risk on households' portfolio composition. A portfolio choice model with both housing and nonhousing consumption is developed to demonstrate that the optimal holding of the risky assets is additionally motivated by households' hedging incentives against unfavorable housing price shocks. Such motive is dependent on location and household lifecycle: it is stronger in places with less elastic housing supply and for young households who are on the rising path of their lifecycle housing consumption profile. Data from recent waves of PSID in the US provide empirical support that that households living in metropolitan areas with less elastic housing supply invest a relatively larger fraction of their financial wealth in risky assets (stocks), and this effect is more pronounced for the young households. These results suggest that financial asset provides important means for households to hedge against housing consumption risk, in addition to the means provided by homeownership adjustment shown in the extant literature.

The third essay is motivated by the work of Wei *et al.* (2011), which shows that the substantial increase in household saving in China since late 1990s may have to do with a rising male-female sex ratio. They find a higher sex ratio in a region in China makes parents with a young son save more for the son's expenses, such as wedding, education and housing, to help the son compete in local marriage market. Two hypotheses are examined in this essay. First, marriage market competition makes young males spend more where the sex ratio is higher – they may do so with financial support from their parents. Second, young males in high sex ratio regions would also work harder, so that their earning would rise faster, in order to pay back their parents in the future. A large dataset of credit card account information of individuals across 31 provinces in China is employed to test these hypotheses. It is found that an additional percentage point in regional sex ratio of age 20 to 34 in 2005 is associated with two to three percent higher credit card balance for males in this cohort but not for females. In addition, young males' age profile of income is steeper in provinces with higher sex ratio. These findings are consistent with the proposed hypotheses and suggest that the rising sex ratio in China may also have contributed to China's high GDP growth through competitive consumer spending and labor supply by young males.

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Chapter 1. Introduction

Three essays in this thesis explore the connection between household consumption/finance choice and macroeconomy performance. In particular, essay one and essay two examine how housing as a consumption good and housing supply elasticity could affect asset risk premia and household asset allocation respectively, and essay three investigates how sex ratio imbalance in China affects the consumption and labor supply behavior of young males. This chapter provides a broad context of these studies, where the research motivation and intended contribution are highlighted.

1.1. Overview of the Research

Housing is not only the dominant wealth component of most households, but also the major component in the household's consumption basket. Given the dual role of housing as both consumption goods and asset, the finance literate has increasing recognized that housing play an important role in affecting asset risk premia and household's portfolio choice. For instance, recent studies has shown that considering the unique features of housing help us to understand the determinants of asset risk premia, e.g., Grossman *et al.* (1990) and Flavin *et al.* (2008) pay particular attention to the transaction cost of housing, Yogo (2006) and Piazzesi *et al.* (2007) focus on the nonseparability of nondurable consumption and durable (housing) consumption, and Lustig *et al.* (2005) emphasize the role of housing as collateral to transmit shocks in housing market to risk premia.

However, these studies normally follow the exchange economy setting of Lucas (1978) where the supply of asset is fixed. In the portfolio choice literature, Brueckner (1997) pioneers works on examining how housing as both a consumption good and an asset can affect asset allocation[1]. Although studies belonging to this strand of literature differ in their particular focuses, they are generally confined in either the mean-variance framework of Markowitz (1952) or the life-cycle choice model of Samuelson (1969) with exogenous specifications of return and price processes, meaning that the feedback from the investment demand on the supply of asset is absent.

While the extant studies mostly focus on the demand side of housing, the general equilibrium effects rising from the supply side of housing are largely overlooked at the present. Why the effects arising from the supply of housing is important for our understanding of the role of housing in shaping asset risk premia and influencing household asset allocation? First, the supply of housing is the other side of market forces that clear the housing market and determine the price of housing. Importantly, it has been found that the price elasticity of housing supply, which is one of the indicators measuring the supply condition of housing market, is crucial in affecting housing price level, volatility, persistence of housing market cycles (Glaeser *et al.* (2006), Glaeser *et al.* (2008), Huang *et al.* (2012), and Paciorek (2013)). Therefore, the processes of asset return and price, and the

[1] See also, e.g., Cocco (2004), Englund *et al.* (2002), Flavin *et al.* (2002), Flavin *et al.* (2011), Hu (2005), Iacoviello *et al.* (2003), Quigley (2006), Yamashita (2003), and Yao *et al.* (2005).

correlations among them, which are purely exogenous but critical in the Lucas' "tree" model and portfolio choice models aforementioned, depend on the supply condition of housing market via the housing supply elasticity. Second, the housing supply elasticity not only has substantial variation across regions that could be attributable to differences in either physical and geographical constraints or regulatory practices (Glaeser *et al.* (2008), Green *et al.* (2005), Ortalo-Magné *et al.* (2011), Quigley *et al.* (2005), and Saiz (2010)), but may also have secular declining trend due to factors like limit amount of developable land, increasing-restricted man-made regulations on housing development, and population concentration in big cities[2]. The time- and geographic variation of housing supply elasticity, which would result in time- and geographic variation of housing price and the correlation between housing price and asset returns, must have implications on intratemporal and intertemporal tradeoff of household consumption, and thus affect household's demand for risk premia and household's portfolio composition.

How does the housing supply elasticity affect asset risk premia and household's asset allocation? These questions are not fully addressed in the literature but the answers are important because: (1) asset risk premia, which are not observable, are fundamental in corporate finance and household finance; with the relative ease

[2] Although time-series estimations of housing supply elasticity are hardly attainable, Sinai (2010) provides circumstantial evidence of declining elasticities in the U.S. He reports that average elasticity of housing supply, measured as the average ratio of the average house price growth over the prior 20 years to the average housing unit growth in each MSA, range from 0.77 over the 1950–1970 period to 0.99 over 1980–2000, indicating "that housing supply in the United States became more inelastic, or less price responsive, over this period" (see Table 1 therein).

of measuring the housing supply elasticity, understanding its relation to asset risk premium would help to forecast asset risk premia; (2) the efficiency of consumption smoothing directly depends on the optimality of asset allocation; understanding how the portfolio choice should adjust according to local housing supply elasticity would help to allocate asset optimally and achieve the goal of smoothing consumption.

The research gap and the importance of answering the questions mentioned above motive the first two essays of the thesis. Particularly, the first essay asks and answers how the housing supply elasticity affects asset risk premia, and the second essays asks and answers how the housing supply elasticity affects the household portfolio composition.

In the first essay, two important theoretical implications are found based on a two-sector general equilibrium model with housing supply: First, different from an exchange model where the presence of housing consumption elevates the equity risk premium through an independent composition risk (Piazzesi *et al.* (2007)), housing consumption in a production economy introduces a substitution benefit and hedge demand for stocks that lower the equity risk premium. Second, because the housing price is more volatile if the housing supply is less elastic, lower housing supply elasticity reduces the equity risk premium via enhanced intratemporal substitution effect but increases the housing risk premium via elevated consumption risk. Empirical analyses using land share of home values as proxy for housing supply elasticity support the model predictions: a lower housing supply elasticity predicts lower excess stock returns but higher excess housing returns, especially in the long-horizon (6-12 quarters) return forecasts.

The second essay addresses the question about the link between housing supply elasticity and household portfolio composition. With a portfolio choice model containing both nonhousing and housing consumption, it demonstrates that lower housing supply elasticity causes households investing more in stocks for the purpose of hedging housing consumption risk. This model implication is supported by empirical analyses using the recent waves of the Panel Study of Income Dynamics (PSID) in the U.S. and Saiz (2010)'s measure of housing supply elasticities in 269 MSAs. It is found that households living in MSAs with less elastic housing supply indeed invest a relatively larger fraction of their financial wealth in risky assets (stocks). In addition, young households on the rising path of housing consumption profile especially do so because they are facing more housing consumption risk.

The background and motivation for the third essay are different from the first two essays. It is under the context of rising sex ratio imbalance in recent decades in China and motived by the work of Wei *et al.* (2011), which argues that the substantial increase in household saving in China since late 1990s is attributable to the rising male-female sex ratio. While Wei *et al.* (2011) shows that parents with a son have competitive saving motive to improve their son' relative standing in the marriage market, the third essay presents the other side of the story by asking whether the young males themselves have competitive spending motive. In addition, the third essay asks whether the sex ratio imbalance impacts males working efforts. Empirical analyses using a large dataset of credit card account information of individuals across 31 provinces in China show that an additional percentage point in regional sex ratio of young adults (20 to 34) is associated with two to three percent higher credit card balance for males in this cohort but not for females. This supports the hypothesis that young males spend in a competitive manner to attract marriage partners. It is also found that young males' age profile of income is steeper in provinces with higher sex ratio. Overall, these findings are supportive to the proposed hypotheses. They suggest that the rising sex ratio in China may also have contributed to China's high GDP growth through competitive consumer spending and labor supply by young males.

1.2. Intended Contribution

This thesis consists of three essays that aim to deepen our understanding of the connections between household consumption/finance choices and macroeconomic performance. The three essays shed light on the relation between housing consumption/supply and asset risk premia, housing consumption/supply and household portfolio choice, and competitive spending and sex ratio imbalance, respectively.

The first essay enhances our understanding of how housing consumption can affect asset risk premia. In particular, it clarifies that adding housing consumption into the consumption basket in a production economy lowers asset risk premia, in contrast to the conclusions in an exchange economy setting of Piazzesi *et al.* (2007). Consistent with the theory, this essay also first shows that lower housing supply elasticity, measured by higher land share of home value, predicts lower equity risk premium but higher housing risk premium. Note that the aggregate housing supply elasticity has been declining due to the gradual exhaustion of developable land and increasing population concentration in big cities, the positive correlation between housing supply elasticity and equity risk premium provides an alternative explanation for the declining equity risk premium observed in recent decades [3]. This understanding of the link between the housing supply elasticity and asset risk premia enables us to have a better vision about the trend of asset risk premia. If we expect the housing supply elasticity continue to trend downward, we would also expect that the equity risk premium would stay on its declining trajectory.

The second essay first establishes the link between the housing supply elasticity and household portfolio choice. It highlights that through the dependence of asset return correlations and housing consumption risk on the housing supply elasticity, the optimal asset allocation is not location-independent. It thus not only

[3] See, e.g., Blanchard *et al.* (1993), Campbell (2008), Claus *et al.* (2001), Fama *et al.* (2002), Jagannathan *et al.* (2001), and Lettau *et al.* (2008).

contributes to the portfolio choice literature by providing explanations for geographic variations in households' asset allocation, but also has important practical implications for the finance service sector. By showing the demand for equity investment to hedge against housing consumption risk, this study suggests a promising demand for financial innovations, such as housing futures and option contracts tied to regional housing price index (Case *et al.* (1993)). As population continues to concentrate in larger and denser metropolitan areas, where housing supply elasticity tends to decrease, such demand is likely to increase.

The third essay advances our understanding of the impact of sex ratio imbalance on social and economic variables. In particular, it demonstrates that sex ratio imbalance results in higher credit card spending and more working efforts of young males, for they have to compete to attract marriage partner. These results are complementary to Wei *et al.* (2011), and are informative for our understanding of China's high GDP growth, for that it may have to do with young males' competitive consumer spending and labor supply behavior induced by sex ratio imbalance.

1.3. Organization of the Thesis

The rest of the thesis is organized as followed: Chapter 2 presents the first essay titled "*Asset Risk Premia in a Production Economy with Housing*", which explores the effects of housing consumption and housing supply elasticity on asset risk premia; Chapter 3 contains the second essay titled "*Portfolio Demand*

and Housing Consumption Risk Hedging: Evidence from Geographic Variations in the Housing Supply Elasticity", which examines how housing supply elasticity has impact on household portfolio composition; Chapter 4 presents the third essay titled "*Competitive Consumption Spending and Labor Supply: Evidence from Regional Differences in Sex Ratio in China*", which investigates the connection between the consumption and labor supply behavior of young males and local sex ratio imbalance. The final chapter, Chapter 5, concludes the thesis, with highlights on limitation of the study and recommendations for future research.

Chapter 2. Asset Risk Premia in a Production Economy with Housing

Abstract: This chapter investigates the covariation between asset risk premia and elasticity of housing supply in a production economy. In contrast to the result based on an exchange-economy model (Piazzesi *et al.* (2007)), the consumption composition risks here are not independent of nonhousing consumption growth so that the presence of intratemporal substitution between nonhousing and housing consumption actually has the effect of mitigating consumption risk. Moreover, a lower housing supply elasticity makes housing price more volatile in response to productivity shocks. It thus increases the housing risk premium via elevated consumption risk but reduces the equity risk premium via enhanced intratemporal substitution effect. Empirical analyses verify the connections between asset risk premia and elasticity of housing supply. Using the land value share of home value as proxy for the housing supply elasticity, I show that lower housing supply elasticity predicts lower excess stock returns but higher excess housing returns, especially in the long-horizon (6-12 quarters) return forecasts.

Key words: risk premium, housing consumption, housing supply elasticity

JEL No.: R1, G1

2.1. Introduction

The housing supply elasticity, as documented in the housing literature, plays important roles in affecting the housing price level and volatility, the persistence of housing market cycles, and urban forms (Fu *et al.* (2010), Glaeser *et al.* (2006), Glaeser *et al.* (2008), Huang *et al.* (2012), and Paciorek (2013)). However, few studies have explored the implications of the housing supply elasticity on asset risk premia. Although several studies in the finance literature has found that housing affects asset risk premia in multiple ways (e.g., Lustig *et al.* (2005), Piazzesi *et al.* (2007)), they are silent about effects arising from the supply side of housing due to the exchange economy setting of Lucas (1978). Because asset risk premia are driving forces of asset price volatilities and the invisible hand behind households' decisions on consumption and saving, and because the housing supply elasticity has not only significant geographic variations but also secular time trend, establishing theoretical linkages between the housing supply elasticity and asset risk premia is of great importance for understanding cross-sectional differences in and time-series profiles of household finance. However, this is an underexplored area in the literature.

To fill the gap, a general equilibrium model with production and housing supply is proposed to examine how asset risk premia change with the housing supply elasticity. Following Piazzesi *et al.* (2007), this paper assumes (1) agents have power utility and consume both nonhousing goods and housing service that are aggregated as a consumption bundle of a CES (constant elasticity of substitution) form, and (2) the intratemporal elasticity of substitution between nonhousing goods and housing service is greater than the intertemporal elasticity of substitution. However, the model in the paper differs from Piazzesi *et al.* (2007) in two important aspects. First, the current model features aggregate productivity shocks in a production economy. This is important because it implies that the growth of nonhousing consumption and its relative quantity to housing consumption are theorectically linked, rather than being independent as in an exchange economy. Second, as the intratemporal elasticity of substitution in a CES function is tied to the price elasticity of demand, this paper assumes an intratemporal elasticity of substitution less than one such that the implied price elasticity of demand for housing, as suggested by empirical evidence, is also less than one. Following the loglinear-lognormal asset pricing approach in Jermann (1998) and Lettau (2003), I first derive expressions for clear understanding of the determinants of asset risk premia in the model, through which I first clarify the role of housing consumption in shaping the asset risk premium and then I analyze the covariation between the housing supply elasticity and asset risk premia.

Similar to Piazzesi *et al.* (2007), it is shown that the risk premium for an asset consists of two components, with the first component reflecting the consumption risk in a model without housing (e.g., CCAPM; called "*consumption risk*" hereafter) and the second component arising from adding housing to the consumption bundle. However, the second component in a production economy setting actually introduces a *substitution benefit* that lowers the equity risk premium, a result opposite to that in an exchange-economy case where the presence of housing introduces an independent consumption composition risk that elevates the equity risk premium (Piazzesi *et al.* (2007)). In particular, because the price of housing service is higher in face of aggregate productivity shocks, the intratemporal substitution of nonhousing consumption for housing consumption results in a higher demand for nonhousing goods. Therefore, if an asset whose payoff is denominated by numeraire goods has higher return in face of shocks, it helps to meet the increased demand for nonhousing goods and thus commands a lower risk premium.

These results suggest that understanding the formation of asset risk premia in fact requires attention not only to the specification of investor preferences, but also to the specification of shocks. An open question about the Housing-CCAPM in Piazzesi *et al.* (2007) is whether it is reasonable to assume shocks to the nonhousing consumption growth and shocks to the expenditure ratio of nonhousing to housing consumption are independent[4]. A positive shock to nonhousing consumption growth implies an increase in equity return and wealth, which in turn increases the demand for housing and thus the housing rent; if nonhousing and housing are substitutes, then depending on whether the intratemporal elasticity is larger or smaller than one, the increase in the housing

[4] The validity of independent sources of uncertainty to nonhousing consumption growth and the expenditure ratio has been questioned by others (see, e.g., Donaldson *et al.* (2007))

rent implies an increase or decrease in expenditure ratio of nonhousing consumption to housing consumption. In short, there are theoretical linkages among nonhousing consumption growth, equity returns, and the expenditure ratio. If a shock affects the nonhousing consumption growth and the expenditure ratio simultaneously, as in the production economy featuring aggregate shocks in the paper, it turns out that the asset pricing implications of the Housing-CCAPM are opposite to those in an exchange setting in which shocks could be independent.

Except for clarifying the role of housing consumption in affecting the asset risk premium, the primary aim of this study is to understand how a decrease in the housing supply elasticity has effect on the asset risk premium. Due to factors such as limited amount of developable land, increasingly-restricted man-made regulations on housing development and population concentration in big cities, the housing supply elasticity may decrease over time. For example, Sinai (2010) provides circumstantial evidence of declining elasticities in the U.S. housing market[5]. Although the secular trend of the housing supply elasticity is endogenous to the factors aforementioned, for simplicity, in the model we consider exogenous changes in the housing supply elasticity. The explicit consideration of housing supply suggests that supply condition in the housing market affects asset risk premia through its impact on the marginal rate of consumption (MRS) and the

[5] Sinai (2010) reports that average elasticity of housing supply, measured as the average ratio of the average house price growth over the prior 20 years to the average housing unit growth in each MSA, range from 0.77 over the 1950–1970 period to 0.99 over 1980–2000, indicating “that housing supply in the United States became more inelastic, or less price responsive, over this period” (see Table 1 therein).

volatility of housing return. For example, when the elasticity of housing supply is low, the agent understands that the housing service cannot be increased in time of need, so they have to substitute nonhousing consumption for housing consumption. This will decrease the MRS, which in turn may result in lower asset risk premia through substitution benefit the asset provides.

However, because the overall effect of changes in the housing supply elasticity on asset risk premia depends on the relative strengths of the consumption risk and the substitution benefit, it is shown that lower housing supply elasticity results in lower equity risk premium but higher housing risk premium. In particular, because inelastic housing supply amplifies housing consumption risk in relation to aggregate productivity shocks, lower housing supply elasticity enhances the substitution benefit provided by equity investment and it thus makes equity investment more attractive and lowers the equity risk premium. For the housing, because its return is more volatile when the housing supply elasticity is low, inelastic housing supply results in an increase in the consumption risk that dominates the increase in the substitution benefit and hence it increases the housing risk premium via elevated consumption risk.

Empirically, the land value share of home value in the U.S. estimated by Davis and Heathcote (2007) is used as proxy for the housing supply elasticity, as the model suggests that higher land value share implies lower housing supply elasticity. Being consistent with the theoretical analysis, it is found that the land value share of home value helps to predict (negatively) excess stock returns and (positively) excess housing returns, especially in the long-horizon (6 – 12 quarters) return forecast. The negative correlation between the land value share (or the positive correlation between the housing supply elasticity) and the equity risk premium suggests the decreasing of the housing supply elasticity in general could potentially account for the declining of the equity risk premium in the U.S. stock market, a fact documented and discussed in the finance literature[6]. As the populations continue to concentrate in large and dense cities and the developable land gradually exhausts, which together would result in a downward trend of the housing supply elasticity, we may expect the equity risk premium to continue to stay in its declining trajectory[7].

In terms of contributions, this paper extends the analysis of Housing-CCAPM developed by Piazzesi *et al.* (2007) in a production economy setting and helps to clarify the role of housing in affecting asset risk premia in a production economy as opposed to an exchange economy. In addition, by showing the theoretical and

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[6] For example, Blanchard *et al.* (1993), Claus *et al.* (2001), Fama *et al.* (2002), and others have argued that the equity risk premium should be substantially lower than the historical average of excess return of 7-8% (see also Campbell (2008), Jagannathan *et al.* (2001), and Lettau *et al.* (2008)). For a recent estimate, the Federal Reserve Bank of Philadelphia's Survey of Professional Forecasters in the first quarter of 2014 found that the predicted median annual-average return on stocks (S&P500) is 6% over the next 10 years. While Treasury bills will return 2.5% annually over the same period, the forcast implies a much lower risk premium of 3.5% compared to the historical average of 6.8%. See http://www.phil.frb.org/research-and-data/real-time-center/survey-of-professional-forecasters/2014/survq114.cfm

[7] Recent studies in the housing literature show that there is an upward trend in land share of home values. For example, Glaeser *et al.* (2005) and Davis and Heathcote (2007) both show that the land value shares were substantially higher in the recent decades than in the 1950s. As we will show in Section 2.2, higher land value share means lower housing supply elasticity.

empirical evidence that the declining equity risk premium is linked to the lower housing supply elasticity, this paper suggests an alternative explanation for the declining equity risk premium.

Although Jaccard (2011) also investigates the implications of housing supply on asset risk premia, this paper differs from Jaccard (2011) in two important aspects. First, I derive analytical expression for the asset risk premium. Compared to *ex post* excess returns calculated from simulated moments in Jaccard (2011), the analytical expression provides clearer understanding about the formation of asset risk premium, the role of housing supply and the conditions under which asset risk premia increases (decreases) with the housing supply elasticity. Second**,** this paper allows for more general utility specification than Jaccard (2011), which specifies a Cobb-Douglas form for the consumption bundle of nonhousing consumption and housing service. It turns out the effects of housing supply elasticity on asset risk premia crucially hinge on specification of utility function.

The rest of the paper is structured as follows: the model is presented in Section 2.2, where the asset risk premium is defined and the role of housing consumption is highlighted; Section 2.3 presents model calibration and discussion of the numerical result, followed by empirical tests of the model implications in Section 2.4; finally, Section 2.5 concludes.

2.2. Model

2.2.1. Asset Risk Premium: A Definition

According to the first order intertemporal optimality condition, an asset with gross return R_t must satisfy the Euler equation:

$$E_t\left(M_{t+1}R_{t+1}\right)=1 \tag{1}$$

where M_{t+1} is the stochastic discount factor in a model and dependent on the specification of utility function. It represents the present value of an extra unit of numeraire consumption tomorrow. As we will see below, a nonseparable utility over nonhousing and housing consumption results in a stochastic discount factor that is a function of both nonhousing consumption growth and changes in the relative quantity of nonhousing consumption to housing consumption.

Let R_t^f be the return on one-period bond, that is, the risk-free rate, equation (1) suggests:

$$R_t^f = \frac{1}{E_t\left(M_{t+1}\right)} \tag{2}$$

Define also the risk premium for an asset with expected return $E_t\left(R_{t+1}\right)$ as the difference between the log of the expected asset return and the log of risk free rate:

$$r_{t+1}^{rp} \equiv \ln\left(E_t\left(R_{t+1}\right)\right) - \ln\left(R_t^f\right) \tag{3}$$

If we use the lower case with time subscript to denote the log of a variable, e.g., $r_t = \ln\left(R_t\right)$ and $m_t = \ln\left(M_t\right)$, and assume that M_t and R_t are jointly lognormal and homoskedastic, the Euler equation (1) suggests that the log of the expected asset return can be expressed as:

$$\ln\left(E_t\left(R_{t+1}\right)\right) = E_t(r_{t+1}) + \frac{1}{2}\operatorname{var}_t(r_{t+1}) = -\left(E_t(m_{t+1}) + \frac{1}{2}\operatorname{var}_t(m_{t+1})\right) - \operatorname{cov}_t(m_{t+1}, r_{t+1}) \tag{4}$$

Using (2), we have the log of risk free rate:

$$r_t^f = -\left(E_t(m_{t+1}) + \frac{1}{2}\operatorname{var}_t(m_{t+1})\right) \tag{5}$$

Equation (3), together with (4) and (5), implies that the asset risk premium is determined by the conditional covariance between the (log of) stochastic discount factor and the (log of) asset return:

$$r_{t+1}^{rp} = -\operatorname{cov}_t(m_{t+1}, r_{t+1}) \tag{6}$$

2.2.2. Nonseparable Utility Specification and Its Asset Pricing Implications

Consider a representative agent model in which the agent has an infinite lifetime and a preference of the standard form over the aggregate consumption $G(C_t, H_t)$:

$$E\left(\sum_{t=0}^{\infty} \beta^t U\left(G\left(C_t, H_t\right)\right)\right) \tag{7}$$

where β is the subjective time discount factor, and $U\left(G\left(C_t, H_t\right)\right) = \frac{\left(G\left(C_t, H_t\right)\right)^{1-\gamma}}{1-\gamma}$ is the period power utility function with γ being the relative risk aversion, or equivalently, with $\sigma = 1/\gamma$ being the intertemporal elasticity of substitution of the aggregate consumption. Following Piazzesi *et al.* (2007), assume the aggregate consumption $G(C_t, H_t)$ is a quantity index that aggregates nonhousing consumption C_t and housing service H_t, and it has the form of constant elasticity of substitution (CES) [8]:

$$G\left(C_t, H_t\right) = \left(\omega C_t^{1-1/\varepsilon} + \left(1-\omega\right) H_t^{1-1/\varepsilon}\right)^{\frac{\varepsilon}{\varepsilon-1}} \tag{8}$$

[8] In what follows, the nonhousing consumption is treated as numeraire consumption, and the terms "nonhousing" and "numeraire" are used interchangeably.

where ω is the share parameter for non-housing consumption, and ε represents the intraperiod elasticity of substitution between non-housing and housing consumption.

The CES preference over nonhousing and housing consumption implies that the marginal utility of nonhousing consumption is a function of nonhousing consumption as well as its relative quantity to housing consumption. As shown in Piazzesi *et al.* (2007), it follows that the stochastic discount factor in the model depends on both the growth of nonhousing consumption and the growth of its relative quantity to housing consumption:

$$M_{t+1} \equiv \beta \left(\frac{C_{t+1}}{C_t} \right)^{-\gamma} \left(\frac{\omega + (1-\omega)\left(\frac{C_{t+1}}{H_{t+1}} \right)^{\frac{1}{\varepsilon}-1}}{\omega + (1-\omega)\left(\frac{C_t}{H_t} \right)^{\frac{1}{\varepsilon}-1}} \right)^{\frac{(\varepsilon-\sigma)}{\sigma(1-\varepsilon)}} \tag{9}$$

Taking the log approximation of (9) around $c_t - h_t = \ln\left(\frac{C_t}{H_t} \right)$, we can obtain the following expression for the log of the stochastic discount factor[9]:

$$m_{t+1} = \ln(\beta) - \gamma(c_{t+1} - c_t) + \kappa_t \left(c_{t+1} - h_{t+1} - (c_t - h_t) \right) \tag{10}$$

[9] See Appendix A at the end of the thesis for the derivation.

where $\kappa_t = \frac{(\varepsilon-\sigma)}{\sigma\varepsilon}\left(1+\left(\frac{\omega}{1-\omega}\right)^{\varepsilon}\left(D_t^h\right)^{\varepsilon-1}\right)^{-1}$ with $D_t^h = \frac{1-\omega}{\omega}\left(\frac{H_t}{C_t}\right)^{-\frac{1}{\varepsilon}}$ being the implicit price of housing service at time t implied by the intratemporal choice. In this paper, I follow Piazzesi *et al.* (2007) and assume the *intratemporal* elasticity of substitution is greater than the *intertemporal* elasticity of substitution ($\varepsilon > \sigma = \frac{1}{\gamma}$) so that $\kappa_t > 0$ [10]. As can be seen from equation (9) or equation (10), the value of the stochastic discount factor (or pricing kernel) is high if nonhousing consumption tomorrow is low. In addition, the present value of extra nonhousing consumption is high if the relative quantity of nonhousing consumption to housing service tomorrow is high.

Substitute (10) into (6), we have the expression for the asset risk premium in the model with housing consumption:

$$r_{t+1}^{rp} = \gamma Cov_t\left(\Delta c_{t+1}, r_{t+1}\right) - \kappa_t Cov_t\left(\Delta\left(c_{t+1} - h_{t+1}\right), r_{t+1}\right) \tag{11}$$

Equation (11) suggests that, in a model with housing consumption, the asset risk premium consists of two components. The first component reflects the *consumption risk* in the conventional CCAPM. If the return of an asset is high

[10] As argued by Piazzesi *et al.* (2007), the assumption that elasticity of intratemporal substitution is greater than elasticity of intertemporal substitution ($\varepsilon > \sigma$) means that the agent is more willing to substitute housing and nonhousing consumption within a period than he is to substitute the overall consumption bundles between periods. This assumption is in line with observations that households are prone to consumption substitution within a certain period. Empirically, Pakoš (2011) finds the intratemporal substitutability (0.18) between nondurable goods and service flow from the stock of durable goods is greater than the intertemporal substitutability (0.04).

(low) when the nonhousing consumption is high (low), then agents require a positive risk premium to hold this asset because it fails to hedge against the consumption risk. The second component reflects how the housing could play roles in affecting asset risk premium. It depends on how the relative quantity of nonhousing consumption to housing consumption is correlated with asset returns. Given that $\kappa_t > 0$, incorporating housing consumption into a consumption-based model would lower the risk premium of an asset if its return is positively correlated with the relative quantity $c_{t+1} - h_{t+1}$.

We can further simplify equation (11) to have a better understanding of the determinants of asset risk premia. The linear approximation approach in the perturbation theory suggests that the logs of endogenous variables in a model can be written as linear functions of the logs of the state variables, with coefficients in the front of the state variables being interpreted as the elasticities of endogenous variables with respect to changes in state variables (see, e.g., Campbell (1994), Judd (1996), King *et al.* (1988), Lettau (2003), Schmitt-Grohe *et al.* (2004), and Uhlig (1998)). In fact, these functions are conventionally called “policy functions”. For example, the policy function for the nonhousing consumption at time $t+1$ can be written as:

$$c_{t+1} = c + \eta_{c,1} s_{1,t} + \eta_{c,2} s_{2,t} + \ldots + \eta_{c,k} s_{k,t} + \eta_{c,\epsilon} \epsilon_{t+1} \qquad (12)$$

where the constant in the linear function (c) denotes the steady state value of nonhousing consumption (in log); $\eta_{x,k}$ denotes the elasticity of endogenous variable x with respect to the k^{th} state variables $s_{k,t}$ in the model (more precisely, the log deviation of the state variable from its steady state), and $\eta_{x,\epsilon}$ denotes the elasticity of endogenous variable x with respect to the shock ϵ_{t+1} [11]. The elasticities of endogenous variables reflect how agents optimally react to changes in state variables and shocks. For instance, if $\eta_{c,\epsilon} = 0.05$, then the numeraire consumption at time $t+1$ is 5 percent above its steady state level in response to 1 unit of shock. Similarly, the housing consumption and the asset return at time $t+1$ can be written as:

$$h_{t+1} = h + \eta_{h,1}s_{1,t} + \eta_{h,2}s_{2,t} + ... + \eta_{h,k}s_{k,t} + \eta_{h,\epsilon}\epsilon_{t+1} \tag{13}$$

$$r_{t+1} = r + \eta_{r,1}s_{1,t} + \eta_{r,2}s_{2,t} + ... + \eta_{r,k}s_{k,t} + \eta_{r,\epsilon}\epsilon_{t+1} \tag{14}$$

Substituting (12), (13), and (14) into (11), we can express the asset risk premium in terms of these elasticities and the variance of the underlying shock σ_ϵ^2:

$$r_{t+1}^{rp} = \gamma\eta_{c,\epsilon}\eta_{r,\epsilon}\sigma_\epsilon^2 - \kappa_t(\eta_{c,\epsilon} - \eta_{h,\epsilon})\eta_{r,\epsilon}\sigma_\epsilon^2 \tag{15}$$

[11] We will discuss the state variables and the shock of the general equilibrium model in this paper when we complete the model with production sectors and the supply of housing service.

Similar to equation (11), equation (15) suggests that the asset risk premium has two components and depends on the responsiveness of consumption and asset return to the shock, namely, the elasticities of nonhousing consumption and housing consumption with respect to the shock ($\eta_{c,\epsilon}$ and $\eta_{h,\epsilon}$) and the elasticity of the asset return with respect to the shock ($\eta_{r,\epsilon}$). If an asset return has an elasticity of the same sign with the elasticity of numeraire consumption, then the first term in (15) is positive, reflecting the consumption risk in the CCAPM. The second component $-\kappa_t(\eta_{c,\epsilon}-\eta_{h,\epsilon})\eta_{r,\epsilon}\sigma_{\epsilon}^{2}$, again, is related to the housing consumption. If the period consumption utility is separable in nonhousing and housing ($\varepsilon=\sigma={}^{1}\!/_{\gamma}$) such that $\kappa_t=0$, the second component is zero and the consumption risk is sufficient to determine the asset risk premium. However, assuming that the intratemporal elasticity of substitution is greater than the intertemporal elasticity of substitution ($\varepsilon>\sigma={}^{1}\!/_{\gamma}$) such that κ_t is positive, we note that the second component is negative (positive) if the asset return is positively (negatively) correlated with the relative quantity of nonhousing consumption to housing consumption ($\eta_{c,\epsilon}-\eta_{h,\epsilon}$). If an asset return reacts to a shock in the same direction as the numeraire consumption (e.g., $\eta_{c,\epsilon}>0$ and $\eta_{r,\epsilon}>0$), but at the same time housing consumption responds more than nonhousing consumption to the shock such that $\eta_{c,\epsilon}-\eta_{h,\epsilon}<0$, then the second component is positive, reflecting the "composition risk" proposed in Piazzesi *et al.*

(2007). In this case, incorporating housing into the model increases the asset risk premium. However, if the numeraire consumption is more responsive to shocks such that $\eta_{c,\epsilon} - \eta_{h,\epsilon} > 0$, then the second component is negative and considering housing in the model lowers the asset risk premium. We will return to this point when we discuss the numerical results in Section 2.3.

Except for clarifying the role of housing consumption in affecting the asset risk premium in a model featuring aggregate shocks, the primary objective of the paper is to understand how the housing supply elasticity affects asset risk premia. More explicitly, the question is how the elasticities of the endogenous variables of interest in equation (15) and thus the asset risk premia change with the housing supply elasticity. To accomplish this, I next add production sectors and the supply of housing service to complete the model.

2.2.3. Complete the Model with Production Sectors and the Supply of Housing Service

Assume the production sectors are composed of two sectors: the non-housing sector and the housing sector. The output in the nonhousing sector can be divided between nonhousing consumption and capital investment, while housing service is produced by the housing sector. The production technology in the non-housing sector is represented by the AK production function:

$$Y_t = A_t \left(K_{t-1}^{b} \right)^{\alpha}, \quad 0 < \alpha < 1 \tag{16}$$

where Y_t is the gross production in the non-housing sector that can be used for both consumption and investment; K_{t-1}^b is the production capital in the nonhousing sector and it is predetermined at the period $t-1$; α determines the output elasticities of K_{t-1}^b, and A_t is the stochastic productivity level. It follows that the gross return on production capital is:

$$R_t^k = \alpha A_t \left(K_{t-1}^b\right)^{\alpha-1} + 1 - \delta_k \tag{17}$$

where δ_k is the depreciation rate of production capital.

Assume the productivity level is the only aggregate uncertainty in the economy and it follows an AR(1) process:

$$\ln\left(A_t\right) = \rho_a \ln\left(A_{t-1}\right) + \epsilon_t, \quad \epsilon_t \sim N(0, \sigma_\epsilon^2) \tag{18}$$

where ρ_a measures the persistence of productivity process and σ_ϵ^2 is the variance of the technology shock.

Furthermore, assume that there is one unit of land in the economy, which is not depreciable and is combined with the production capital (structure) in the housing sector K_{t-1}^h to produce housing service[12]:

$$H_t = \left(K_{t-1}^h\right)^{\varphi} \tag{19}$$

where $0 < \varphi < 1$ determines the structure share in the cost of housing service, or equivalently, $1-\varphi$ presents the share of land value in the cost of housing service. In fact, φ also governs the responsiveness of the supply of housing service to changes in the price of housing service D_t^h and the gross return on capital R_t^k. To see this, note the optimal allocation of capital between the nonhousing sector and the housing sector requires that the marginal product of capital are equalized in these two sectors:

$$R_t^k - 1 + \delta_k = \varphi\left(K_{t-1}^h\right)^{\varphi-1} D_t^h \tag{20}$$

Equation (20) implies a housing supply function, with the price of housing service and the gross return on capital as its arguments:

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[12] The assumption that land supply is fixed at one unit is not unrealistic if we consider that the total amount of land in a country is normally unchanged overtime. Because the housing service is generated by land and the structure attached to it, this assumption also implies that one unit of land is required as input for producing one unit of housing service but the marginal product of land increases as φ decreases. As the land generate revenue $(1-\varphi)H_t D_t^h$, the land price P_t^l satisfies $P_t^l = E_t(M_{t+1}(P_{t+1}^l + (1-\varphi)H_{t+1}D_{t+1}^h))$ and can be solved from this equation. However, I do not discuss the property of land price in this model.

$$\ln\left(H_t\right) = \frac{\varphi}{1-\varphi}\ln\left(\varphi\right) + \frac{\varphi}{1-\varphi}\ln\left(D_t^h\right) + \frac{\varphi}{\varphi-1}\ln\left(R_t^k - 1 + \delta_k\right) \tag{21}$$

By definition and construction, the housing supply elasticity as suggested in (21) is:

$$\eta = \frac{\varphi}{1-\varphi} \tag{22}$$

Note that because $0 < \varphi < 1$, $\eta \in (0, +\infty)$. Importantly, a smaller φ implies both higher share of land value in the cost of housing service but lower housing supply elasticity – a negative correlation that is supported by empirical evidence.

Given (16), (18), and (19), the representative agent's problem is to optimally choose nonhousing consumption C_t, and allocate the production capital in the nonhousing sector K_t^b and the production capital in the housing sector K_t^h, subject to the budget constraint (23), such that the expected consumption utility is maximized:

$$\max_{\{C_t, K_t^b, K_t^h\}_{t=0}^{\infty}} E\left(\sum_{t=0}^{\infty} \beta^t U\left(G_t\left(C_t, H_t\right)\right)\right)$$

s.b.t.:

$$C_t + K_t^b + K_t^h = Y_t + \left(1-\delta\right)\left(K_{t-1}^b + K_{t-1}^h\right) \tag{23}$$

The first-order conditions of the maximization problem indicate that:

$$U_c^{'}\left(G\left(C_t, H_t\right)\right) = \beta E_t\left[U_c^{'}\left(G\left(C_{t+1}, H_{t+1}\right)\right)\left(\alpha A_{t+1}\left(K_t^b\right)^{\alpha-1} + 1 - \delta_k\right)\right] \quad (24)$$

$$U_c^{'}\left(G\left(C_t, H_t\right)\right) = \beta E_t\left[\varphi U_h^{'}\left(G\left(C_{t+1}, H_{t+1}\right)\right)\left(K_t^h\right)^{\varphi-1} + U_c^{'}\left(C_{t+1}, H_{t+1}\right)\left(1 - \delta_k\right)\right]$$

(25)

where $U_c^{'}\left(G\left(C,H\right)\right)$ and $U_h^{'}\left(G\left(C,H\right)\right)$ are the first order derivatives of $U(G\left(C,H\right))$ with respect to its first and second argument. Equation (24) and (25) represent the intertemporal tradeoff of consumption and savings and the optimal allocation of production capital in the two sectors, respectively. Together with equation (16) - (19), and equation (23), they define a general equilibrium model under rational expectation, with the production capital K_{t-1}^b and K_{t-1}^h being the endogenous state variables, the productivity level A_t being the exogenous state variable, and all of the rest being endogenous variables.

In the next section, I calibrate the model and solve the elasticities in equation (15) numerically using Dynare[13]. This will allow us to understand the relative responsiveness of consumption and asset returns to shocks and how the housing consumption plays a role in affecting the asset risk premium. Importantly, as φ is tied to the housing supply elasticity η, we experiment with different values of φ to examine the impact of housing supply (η) on the responses of endogenous variables and assert risk premia to the technology shocks.

2.3. Numerical Analysis

2.3.1. Calibration

The model is calibrated to quarterly data and a summary of the parameter calibration is provided in Table 2.1. Most of the parameter values are standard in the literature. The subjective discount factor β is set to 0.98. Following Lustig *et al.* (2010), I consider the case of intratemporal elasticity of substation less than one ($\varepsilon = 0.8$), which suggests that the price elasticity of demand for housing service is also less than one. The nonhousing expenditure share is set at $\omega = 0.8$

[13] Dynare is a software platform specialized in solving dynamic stochastic general equilibrium (DSGE) models. Dynare first finds the deterministic steady states of the nonlinear DSGE model starting from initial guesses provided by the user. Once the steady states are found, Dynare will check the stability conditions of the model. If the Blanchard-Kahn stability conditions (Blanchard *et al.* (1980)) are satisfied so that there is a unique and stable equilibrium in the neighborhood of the deterministic steady state, Dynare uses a pure perturbation approach as in Schmitt-Grohe *et al.* (2004) by default to solve for policy functions and state transition questions of the model. As a user-friendly tool, Dynare is not only widely used by various public bodies and private financial institutions for performing policy analysis and forecasting, but also popular in economic research and teaching. See Adjemian *et al.* (2011) and http://www.dynare.org/. For some typical applications in the housing/real estate literature, see, e.g., Ghent (2012), Iacoviello *et al.* (2010), Jaccard (2011), and Liu *et al.* (2013).

according to Davis and Ortalo-Magne (2007), who find that the MSA-average of the housing expenditure is at 24 percent with little variation. For the degree of risk aversion, we set $\gamma = 5$ as the low risk aversion benchmark. A high risk aversion case with $\gamma = 12$ is also investigated to explore the sensitive of the results with respect to different degree of risk aversion. The capital share in the production of numeraire goods and its depreciation rate are set at $\alpha = 0.33$ and $\delta_k = 0.025$ (10% at annual rate), as in Campbell (1994). For the capital share in the production of housing service (φ), I experiment with a range of parameter values from 0.4 to 0.8, implying a range of housing supply elasticity from 0.67 to 4. These allow us to examine how the elasticities of endogenous variables in equation (15) and thus how asset risk premia change with the housing supply elasticity. Finally, I consider a fairly persistent technology shock $\rho_a = 0.95$.

Table 2.1 Parameterization

Parameter		Value
Preference		
β	Subjective time discount factor	0.98
ε	Elasticity of intratemporal substitution	0.8
ω	Non-housing expenditure share	0.8
γ	Constant relative risk aversion	5 and 12
Technology		
α	Capital share in the production of numeraire goods	0.33
δ_k	Depreciation rate of production capital	0.025
φ	Capital share in the production of housing service	0.4, 0.6, and 0.8
Exogenous shock		
ρ_a	Persistence in productivity shock	0.95

2.3.2. Numerical Results

Table 2.2 presents the numerical results for the elasticity of nonhousing consumption $\eta_{c,\epsilon}$, the elasticity of housing consumption $\eta_{h,\epsilon}$, the elasticity of return on production capital $\eta_{rk,\epsilon}$, and the elasticity of housing return $\eta_{rh,\epsilon}$ for a variety of parameters of the general equilibrium model discussed above[14]. The mean values of the equity risk premium (*erp*) and the housing risk premium (*hrp*) calculated according to equation (15) are also reported.

Table 2.2 Numerical results

(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
γ	φ	η	$\eta_{c,\epsilon}$	$\eta_{h,\epsilon}$	$\eta_{rk,\epsilon}$	$\eta_{rh,\epsilon}$	D^h	κ	*erp*	*hrp*
5	0.8	4	0.4102	0	0.0423	0.0918	0.0839	0.6275	3.1058	6.7422
	0.6	1.5	0.4088	0	0.0423	0.1684	0.1557	0.6948	3.0474	12.1424
	0.4	0.67	0.4083	0	0.0423	0.2573	0.2856	0.8018	2.9931	18.2140
12	0.8	4	0.4062	0	0.0423	0.0955	0.0839	1.7989	7.1756	16.2089
	0.6	1.5	0.4040	0	0.0423	0.1769	0.1557	1.9918	7.0006	29.2937
	0.4	0.67	0.4016	0	0.0423	0.2720	0.2856	2.2985	6.8174	43.8633

Notes: The table reports, for different combination of risk aversion (γ) and the capital share in the production of housing service (φ), the elasticity of nonhousing consumption ($\eta_{c,\epsilon}$), the elasticity of housing consumption ($\eta_{h,\epsilon}$), the elasticity of equity return ($\eta_{rk,\epsilon}$), and the elasticity of housing return ($\eta_{rh,\epsilon}$) with respect to shocks, as well as the mean of equity risk premium (*erp*) and the mean of housing risk premium (*hrp*) calculated according to equation (15). η denotes the housing supply elasticity implied by φ (see equation (21) and (22)). D^h is the steady state value of the price of housing service, and κ is the steady state value of κ_t in equation (15), which is an increasing function of D^h.

[14] In the paper, the housing service is generated by the structure (production capital in the housing sector) and the (fixed and one unit of) land attached to it. Therefore, housing as an asset is not explicitly modeled as an asset in the paper but its price P_t^h can be understood as the present value of the price of housing service and it satisfies $P_t^h - D_t^h = E_t(M_{t+1}(1-\delta_k)\, P_{t+1}^h)$. The housing return in the model is defined as $R_t^h = \frac{(1-\delta_k)\, P_t^h}{P_{t-1}^h - D_{t-1}^h}$. Return on housing differs from return on production capital, which is understood as stock returns in the paper.

We can observe from Table 2.2 that, regardless of the degree of risk aversion and the housing supply elasticity, all of $\eta_{c,\epsilon} - \eta_{h,\epsilon}$, $\eta_{rk,\epsilon}$ and $\eta_{rh,\epsilon}$ are positive so that the second term in equation (15) is negative[15]. Importantly, the positive sign of $\eta_{c,\epsilon} - \eta_{h,\epsilon}$ in the model is not artificial. Consumer theory also suggests that the numeraire consumption should be more responsive to shocks such that $\eta_{c,\epsilon} - \eta_{h,\epsilon} > 0$. To understand this, note that if the relative price of housing, e.g., the housing rent, would not change in response to a shock, then a change in total consumption expenditure resulting from optimal consumption-saving decision after the shock would not alter the relative quantity of consumption goods, so $\eta_{c,\epsilon} - \eta_{h,\epsilon} = 0$ and the second term in equation (15) collapse to zero. However, because the housing rent and the asset return are positively correlated, a positive sign of the elasticity of asset return implies that the housing rent will increase in face of the shock, which in turn will result in a decline in the relative quantity of housing consumption due to intratemporal substitution so that $\eta_{c,\epsilon} - \eta_{h,\epsilon} > 0$. Therefore, both numerical results and the theoretical reasoning suggest that the sign of $\eta_{c,\epsilon} - \eta_{h,\epsilon}$ is positive.

[15] Note that by construction, the production capital in the housing sector is predetermined, so the elasticity of housing consumption is essentially zero irrespective of the degree of risk aversion and the housing supply elasticity. Meanwhile, although the elasticity of nonhousing consumption decreases when the housing supply elasticity is low, the effect is almost negligible. This is because the price of housing service and thus the price of aggregate consumption is high when the housing supply elasticity is low, resulting a smaller consumption quantity adjustment when agents make consumption-saving decisions in response to shocks.

Because of the positive sign of $\eta_{c,\epsilon} - \eta_{h,\epsilon}$, the second term in equation (15), which is negative, reflects a *substitution benefit* provided by an asset whose return is positively correlated with $\eta_{c,\epsilon} - \eta_{h,\epsilon}$, and the substation benefit lowers the asset risk premium. This can also be understood from the view that when the price of housing service is expected to be high or when there is an expected shortfall in housing service, an asset whose payoff is denominated by numeraire goods possesses lower risk premium because agents are induced to hold this asset for the purpose of hedging housing price risk. These results suggest that if a model has the shock to nonhousing consumption growth and the shock to its relative quantity to housing consumption not independent, such as a production-economy setting featuring aggregate productivity shocks we consider here, the *"substitution benefit"* introduced by adding housing to the consumption basket lowers the equity risk premium. This is a conclusion in contrast to Piazzesi *et al.* (2007).

In addition, although lower housing supply elasticity enhances the substitution benefits of an asset, the overall effect depends on how the elasticity of asset return changes with the housing supply elasticity. Column (8) and (9) of Table 2.2 suggest that both the price of housing service D^h and κ in the second term of equation (15), which is an increasing function of D^h, are high when the housing supply elasticity is low. As a result, the substitution benefit becomes more important in determining asset risk premia. For the production capital, because the elasticity of its return is constant (column (6) of Table 2.2), the enhanced substitution effect dominates the consumption risk[16]. Consequently, as shown in column (10) of Table 2.2, the equity risk premium decreases when the housing supply elasticity is low. On the contrary, because the elasticity of housing return with respect to productivity shocks is high when the housing supply elasticity is low, a lower housing supply elasticity elevates the consumption risk of housing returns[17]; as the elevated consumption risk dominates the substitution benefits, the housing risk premium increases when the housing supply elasticity is low, as shown in column (11) of Table 2.2.

2.4. Testable Implications

2.4.1. Predicting Excess Returns with the Land Value Share of Home Value

The analyses in Section 2.3 suggest that lower housing supply elasticity imply lower equity risk premium because of enhanced substitution effect, but it also implies higher housing risk premium because of dominated consumption risk. Although a time series measure of the housing supply elasticity is hardly obtainable, as the model shows that the land value share of housing cost is tied to

[16] The constant elasticity of equity return can be seen from the linear approximation of the equity return based on (17): $r_t^k \approx \frac{\alpha K^b}{R^k}(a_t + (\alpha - 1)k_{t-1}^b)$, where the small letters are the log of corresponding variables and K^b and R^k are the steady state values of corresponding variables. The elasticity of the equity return with respect to the productivity is determined by $\frac{\alpha K^b}{R^k}$ and not affected by risk aversion parameter and the housing supply elasticity.

[17] This is consistent with Kiyotaki *et al.* (2011), who argue that the housing price is more sensitive to shocks in general productivity growth rates in an economy where the land share of housing value is high.

the housing supply elasticity, we can test the model predictions using the time series data on land value share of home value[18].

To see whether the land value share is negatively correlated with the equity risk premium and positively correlated with the housing risk premium, I regress the excess stock (or housing) returns on the land value share:

$$\sum_{k=1}^{n} excessReturn_{t+k} = \beta_0 + \beta_1 \times ls_t + \Phi Z_t + e_{t+1} \qquad (26)$$

where the dependent variable $\sum_{k=1}^{n} excessReturn_{t+k}$ is the *n*-period accumulative excess return for stocks or housing, with the excess return being defined as the total rate of return of stock or housing minus the prevailing short-term interest rate, and the predicting variable is land value share of home value ls_t. As discussed below, Z_t includes a set of variables that have been found useful in predicting excess returns in the finance literate. A significant β_1 with expected sign will lend support to the model implications. I proceed to discuss the data and construction of the variables before presenting the empirical results.

[18] Because agents require lower equity risk premium and are induced to hold more equity investment when the housing supply elasticity is low, another testable implication is that households living in inelastic places should invest more in stocks. Using the geographic variations in the housing supply elasticity measured by Saiz (2010) to identify households' hedging motive to hold stocks, Lai *et al.* (2014) find that the geographic variation in the housing supply elasticity helps to explain portfolio heterogeneity in households' portfolio composition.

2.4.2. Data

I obtain the land value share of home value as the log difference between the aggregate market value of residential land and the aggregate market value of homes using the quarterly estimates of the price of land and housing used for residential purposes in the aggregate U.S. created by Davis and Heathcote (2007)[19]. Housing return is defined as $(Price_t + Rent_t)/\,Price_{t-1}$, and the excess housing return is prepared using the quarterly estimates of housing rents and housing price for the aggregate housing stock in the U.S. provided by Davis *et al.* (2008)[20]. The quarterly finance data, including the stock return, risk-free rate and other controls included in Z_t are from Welch *et al.* (2008), who comprehensively reexamine the performance of variables that the academic literature has suggested to be good predictors of the equity risk premium[21]. Depending on the specifications of the linear regression models, the set of controls Z_t may include the dividend yield (*dy*), the market ratio (*bm*), the net

[19] Davis and Heathcote (2007) constructs both quarterly and annual data on land price and quantity based on data on housing values and estimates of structure costs using price indexes for housing and construction costs. They first benchmark the house price data on an estimate of the value of the stock of housing based on micro data from the 2000 Decennial Census of Housing and 2001 Residential Finance Survey and then extrapolate the benchmark forwards and backwards using either the Macromarkets LLC (formerly Case-Shiller-Weiss) repeat-sales index or the Federal Housing Finance Agency (formerly Office of Federal Housing Enterprise Oversight) repeat-sales index. We use the data constructed with CSW-based price index; using data constructed with FHFA-based price index produces similar results in the analyses. The data is available at https://www.lincolninst.edu/subcenters/land-values/price-and-quantity.asp.

[20] Davis *et al.* (2008) first use micro data from each of the 1960, 1970, 1980, 1990, and 2000 Decennial Census of Housing, annual rents paid for rentals units to estimate annual rents for owner-occupied units, and then divide these estimated rents by the average (self-reported) value of the owner-occupied housing units to produce five benchmark rent-price ratios in each of these years. Rent index from Bureau of Labor Statistics (BLS) and price indexes from either Case-Shiller-Weiss (CSW) or OFHEO are used to interpolate between these benchmark years and to extrapolate beyond 2000 in order to obtain rent value and average housing price. The data is available at https://www.lincolninst.edu/subcenters/land-values/rent-price-ratio.asp

[21] The data is available at http://www.hec.unil.ch/agoyal/

equity expansion (*nits*; a measure of corporate issuing activity; see Boudoukh *et al.* (2007)), the term spread (*tms*, defined as the difference between the long term yield on government bonds and the Treasury-bill), the default yield spread (*dfy*; defined as the difference between BAA and AAA-rated corporate bond yields), the investment to capital ratio (*ik*; see Cochrane (1991)) and the regression residuals obtained from estimating a long-run cointegration relation among consumption, wealth and income (*cay*; see Lettau *et al.* (2001)). Given the availability of data, the final data used in the analysis spans the period from 1975Q1 to 2012Q4.

Table 2.3 presents the summary statistics of the data. According to Table 2.3, the mean of excess stock return is roughly the same as the mean of excess housing return. However, the first order autocorrelation of (excess) housing returns is significant while the counterpart of stock returns is not. These are consistent with general perception that stock returns are highly volatile whereas housing returns appear to be more persistent. The land value share has is mean of 32.21% with a standard deviation of 6.34% over the sample period. Meanwhile, similar to the financial variables contained in Z_t, the first order autocorrelation of land value share is quite high, suggesting that the land value share is highly persistent and may have a secular trend. Nevertheless, it is reasonable to believe that the land value share is stationary in theory, because if the land value share is a random walk, the probability of staying in a finite range of [0, 1] forever is zero, an implication incompatible with the theoretical definition[22].

Table 2.3 Summary statistics (1975Q1 - 2012Q4)

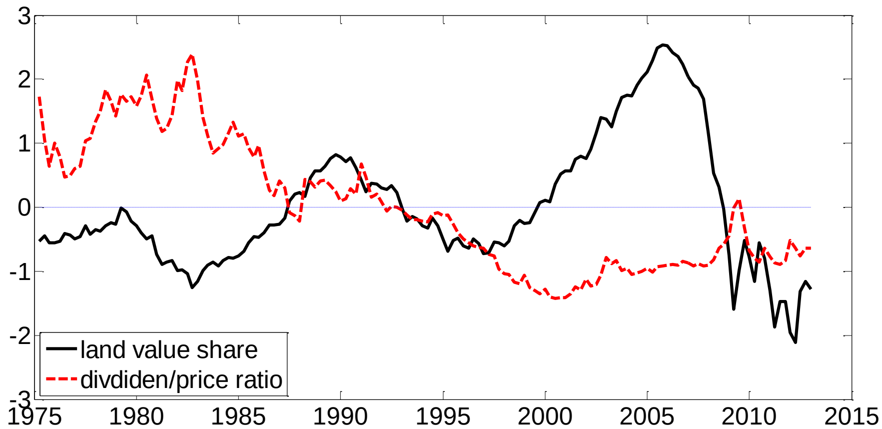
Variable	Mean	Std	Min	Max	1st order autocorrelation
Stock return	0.0539	0.0850	-0.2049	0.2694	0.1095
Housing return	0.0613	0.0237	-0.0323	0.1231	0.6762
Risk free rate	0.0129	0.0085	0.0000	0.0387	0.9482
Excess stock return	0.0409	0.0843	-0.2193	0.2557	0.0842
Excess housing return	0.0483	0.0222	-0.0341	0.1229	0.6269
Land value share	0.3221	0.0634	0.1881	0.4827	0.9814
Dividend yield (*dy*)	0.0303	0.0135	0.0111	0.0625	0.9739
Rent price ratio (*dp*)	0.0474	0.0061	0.0303	0.0562	0.9853
Book to market ratio (*bm*)	0.4870	0.2921	0.1252	1.2016	0.9799
Term spread (*tms*)	0.0216	0.0150	-0.0350	0.0453	0.7996
Default yield spread (*dfy*)	0.0113	0.0048	0.0055	0.0338	0.8493
Net equity expansion (*nits*)	0.0081	0.0200	-0.0532	0.0457	0.9199
Investment to capital ratio (*ik*)	0.0356	0.0038	0.0274	0.0443	0.9767
Consumption, wealth and income co-integration residual (*cay*)	0.0040	0.0216	-0.0432	0.0449	0.9595

As argued by Campbell (2008), the log-linear approximation of dividend price ratio proposed in Campbell *et al.* (1988a, 1988b) suggests that the dividend price ratio is a good proxy for risk premium if the dividend growth is not too predictable. Because there is little evidence for the predictability of dividend growth, we can instead examine the relation between the land value share and dividend price ratio to see whether high land value share is associated with low

[22] Piazzesi *et al.* (2007) apply the same theoretical reasoning to argue for the stationary of nonhousing consumption share (although it is highly persistent too).

equity risk premium. Therefore, as suggestive evidence, Figure 2.1 plots a time series plot of standardized versions of the land value share and the log of dividend to price ratio[23]. It is evident from the figure that the two variables are highly negatively correlated: When the land value share is above its mean, the dividend price ratio is low, suggesting a decline in the equity risk premium; the opposite is true when the land value share is below the mean. This provides statistical evidence for the model predictions in the theoretical sections.

Figure 2.1 Standardized land value share and D/P ratio



Notes: This figure plots the time series of land value share and dividend/price ratio of S&P 500 index. Variables are standardized by subtracting mean and then dividing by standard deviation. Data is quarterly from 1975Q1 to 2012Q4.

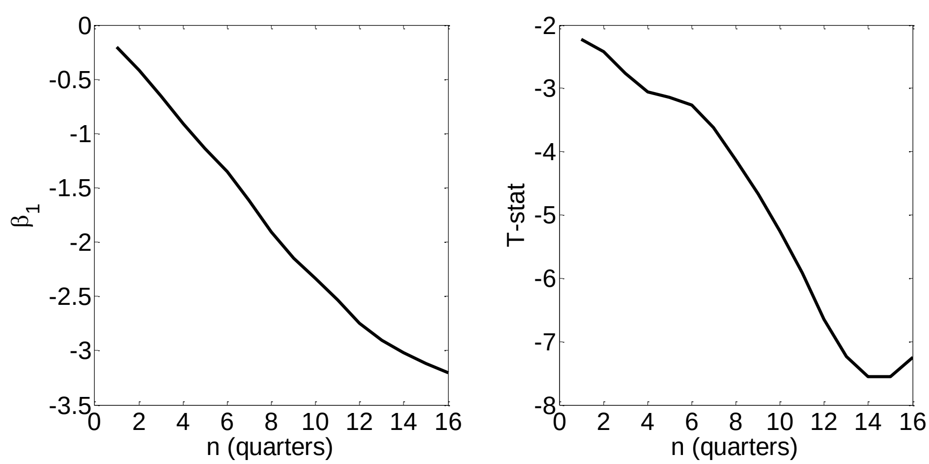
2.4.3. Empirical Results

I first run univariate regressions for the excess stock returns for $n=1, \ldots, 16$ quarters. This means that, for each n, I regress the n-period accumulative excess stock return on a constant and the land value share without including any controls

[23] To standardize a viable, we subtract its mean and then divide it by its standard deviation.

in Z_t. The results are presented in Figure 2.2, with coefficients on the left panel and the corresponding Newey-West *t*-statistics for the coefficients on the right panel. As can be seen, the coefficients are negative, with the Newey-West t-statistics well below -2 for all horizons. This suggests that higher land value share predicts lower equity risk premium, a result consistent with the model prediction. In addition, as *n* increases from 1 to 16, the coefficients increase too and remain significant. While noises may dominate in the short-term return prediction, the increasing predictability of the accumulative excess returns implies that the information contained in the land value share becomes more important for long-horizon forecast.

Figure 2.2 Coefficients and Newey-West *t*-statistics in univariate excess stock return regressions

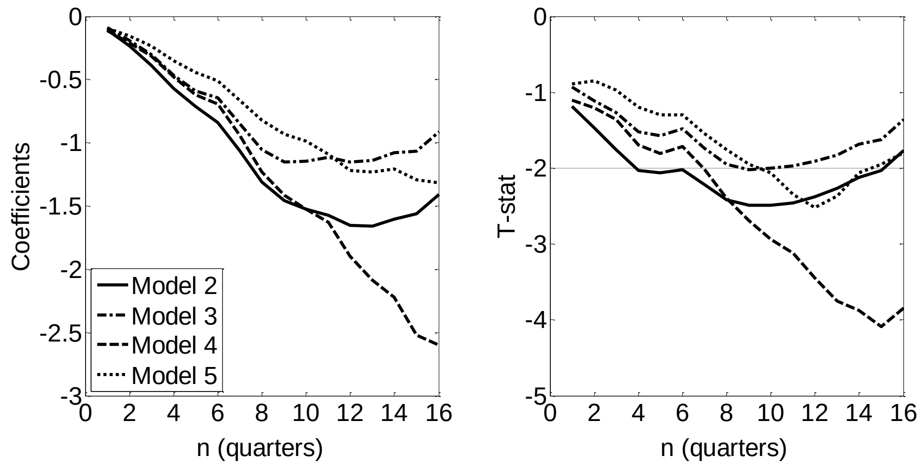


Notes: This figure shows the coefficients on land value share (β_1) and the corresponding Newey-West *t*-statistics in univariate regressions for *n*-period (*n*=1, …, 16) accumulative excess stock returns. The regression model is $\sum_{k=1}^{n} excessReturn_{t+k} = \beta_0 + \beta_1 \times ls_t + e_{t+1}$

To see whether the land value share is simply a proxy for other predictors that have been found useful for predicting equity risk premium, these predictors are gradually added into Z_t to have multivariate regressions. Explicitly, I add dividend yield (*dy*) in model 2; in model 3, three more controls are added, including market ratio (*bm*), term spread (*tms*), default yield spread (*dfy*); model 4 includes another two more controls: net equity expansion (*ntis*) and investment to capital ratio (*ik*); and finally, in addition to all the controls mentioned above, Z_t in model 5 also includes the cointegration error of consumption, wealth and income (*cay*). As our particular interest is the prediction performance of land value share, I first present its coefficient and the corresponding Newey-West *t*-statistic for model 2 to model 5 and $n = 1, \dots, 16$ in Figure 2.3.

Although the order of adding controls is somehow arbitrary, we can see that the coefficients on the land value share (the left panel of Figure 2.3) are consistently negative for all model specifications and all horizons. From the right panel of Figure 2.3, we can see that although the coefficients are marginally significant for short-horizon forecast, they become significant for long-horizon forecast (e.g., when n>6), especially for model 2 and model 4. These results are similar to those in the univariate regressions and lend further support to the model implications.

Figure 2.3 Coefficients and Newey-West *t*-statistics in multivariate excess stock return regressions



Notes: This figure shows the coefficients on land value share (β_1) and the corresponding Newey-West *t*-statistics in multivariate regressions for *n*-period (*n*=1, …, 16) accumulative excess stock returns. The regression model is $\sum_{k=1}^{n} excessReturn_{t+k} = \beta_0 + \beta_1 \times ls_t + \Phi Z_t + e_{t+1}$. I gradually add controls into Z_t in various models. In model 2, Z_t includes only dividend yield (*dy*). In model 3, three more controls are added, including market ratio (*bm*), term spread (*tms*), default yield spread (*dfy*). In model 4, two more controls, net equity expansion (*ntis*) and investment to capital ratio (*ik*), are added. Except for all the controls mentioned above, Z_t in model 5 also includes the cointegration error of consumption, wealth and income (*cay*).

To get a sense of the model specifications and the prediction performance of control variables, Table 2.4 reports regression results for a particular accumulation period n = 8, with the left panel for the regression results of accumulative excess stock returns and the right panel for the regression results of accumulative excess housing returns (discussed below). While the coefficients on other control variables generally have expected sign, it is worth mentioning that the Durbin–Watson (DW) statistics for all model specifications are close to zero, indicating that there are serial correlations in the residuals. However, Augmented Dicker-Fuller (ADF) tests for a unit root of the residuals are all rejected.

Table 2.4 Predicting accumulative excess returns with land value share (accumulative horizon n=8)

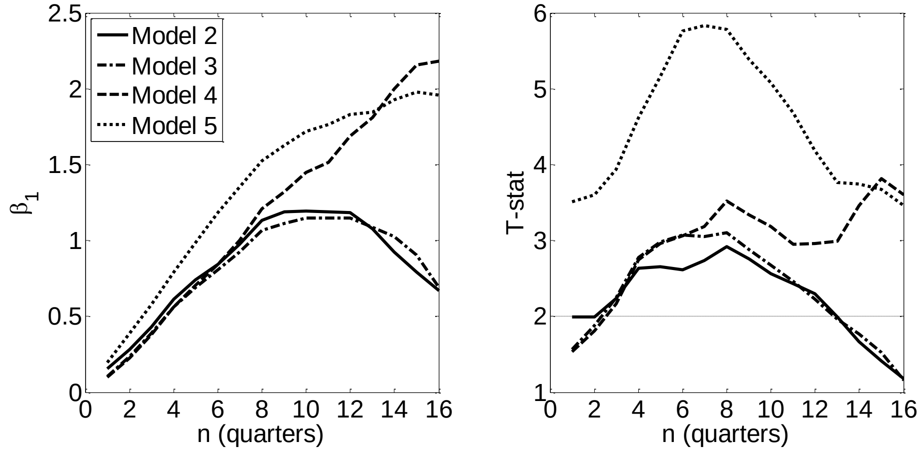
	Excess stock returns				Excess housing returns			
	Model 2	Model 3	Model 4	Model 5	Model 2	Model 3	Model 4	Model 5
Land value share	-1.3088	-1.0531	-1.2322	-0.8224	1.1344	1.0654	1.2098	1.5268
	(-2.4179)	(-1.9464)	(-2.4022)	(-1.7559)	(2.9174)	(3.1013)	(3.5200)	(5.7797)
Dividend yield (*dy*)*	5.0250	12.0767	11.0956	-10.6555	23.0490	22.5881	21.4239	29.8023
	(1.9547)	(2.4970)	(2.4340)	(-1.9028)	(6.0358)	(6.2141)	(6.7176)	(11.0351)
Book to market ratio (*bm*)		-0.2493	-0.2217	0.7469		-0.0325	-0.0319	-0.0965
		(-1.3320)	(-1.1061)	(3.0876)		(-0.7017)	(-0.8287)	(-2.7497)
Term spread (*tms*)		3.9796	2.6337	1.4168		2.3715	3.8133	3.6912
		(2.5270)	(1.5125)	(1.1195)		(2.7600)	(4.3603)	(4.2395)
Default yield spread (*dfy*)		-3.4599	-5.1932	1.0203		-0.0889	3.6998	2.6665
		(-0.5978)	(-0.7963)	(0.2007)		(-0.0410)	(1.5645)	(1.1736)
Net equity expansion (*nits*)			-0.1245	-1.6251			1.0395	0.9241
			(-0.1020)	(-1.8496)			(2.3790)	(2.1141)
Investment to capital ratio (*ik*)			-9.1938	-7.9251			10.9174	8.1404
			(-1.0814)	(-1.0418)			(3.2676)	(2.8479)
cay				6.7530				-2.0895
				(5.9933)				(-4.3066)
Constant	0.5823	0.3588	0.8134	0.7589	-1.0818	-1.0708	-1.5379	-1.8800
	(2.4698)	(1.3324)	(2.1078)	(2.3313)	(-3.8624)	(-4.0846)	(-4.7033)	(-7.6088)
Adj-R^2	0.3023	0.3942	0.4119	0.5579	0.5173	0.6213	0.7001	0.7575
N	144	144	144	144	144	144	144	144
DW	0.3139	0.4883	0.4633	0.3260	0.0481	0.1299	0.3164	0.4206
Unit root of residual	N	N	N	N	N	N	N	N

Notes: This table shows regression results of accumulative excess stock returns and accumulative excess housing returns on a constant, land value share, and other controls included in Z_t: $\sum_{k=1}^{n} excessReturn_{t+k} = \beta_0 + \beta_1 \times ls_t + \Phi Z_t + e_{t+1}$, for a particular n=8. In model 2, Z_t includes only dividend yield (*dy*). In model 3, three more controls are added, including market ratio (*bm*), term spread (*tms*), default yield spread (*dfy*). In model 4, two more controls, net equity expansion (*ntis*) and investment to capital ratio (*ik*), are added. Except for all the controls mentioned above, Z_t in model 5 also includes the cointegration error of consumption, wealth and income (*cay*). We test for a unit root of regression residuals by conducting Augmented Dickey-Fuller (ADF) test with no lagged difference terms. Newey-West *t*-statistics are in parentheses.
*: It is rent price ratio (*rp*) when dependent variable is the accumulative excess housing returns.

I repeat the multivariate regression analyses for the excess housing returns and report the results in Figure 2.4, with the results of a particular accumulation horizon n=8 presented on the right panel of Table 2.4. In contrast to the results in the excess stock return regressions, we can see from the left panel of Figure 2.4 that the coefficients on the land value share are *positive* for all model specifications and all horizons, suggesting that higher land value share predicts higher housing risk premium. This is in line with the model prediction in Section 2.3. The Newey-West t-statistics on the right panel of Figure 2.4 suggests that the performance of the land value share in predicting excess housing returns is better for horizons between 2 and 12, a result similar to that in the excess stock returns. Again, the Durbin–Watson (DW) statistics and results from Augmented Dicker-Fuller (ADF) tests for a unit root suggest that the residuals of excess housing return regressions are autocorrelated but stationary.

Overall, the empirical analyses provide evidence in line with the model implications: on one hand, when the housing supply elasticity is low, or the land value share of home value is high, equity risk premium is low due to the enhanced substitution benefits provided by equity investment; on the other hand, lower housing supply elasticity implies higher housing risk premium because of elevated consumption risk brought by housing investment.

Figure 2.4 Coefficients and Newey-West *t*-statistics in multivariate excess housing return regressions



Notes: This figure shows the coefficients on land value share (β_1) and the corresponding Newey-West *t*-statistics in multivariate regressions for *n*-period (*n*=1, …, 16) accumulative excess housing returns. The regression model is $\sum_{k=1}^{n} excessReturn_{t+k} = \beta_0 + \beta_1 \times ls_t + \Phi Z_t + e_{t+1}$. I gradually add controls into Z_t. In model 2, Z_t includes only dividend yield (*dy*). In model 3, three more controls are added, including market ratio (*bm*), term spread (*tms*), default yield spread (*dfy*). In model 4, two more controls, net equity expansion (*ntis*) and investment to capital ratio (*ik*), are added. Except for all the controls mentioned above, Z_t in model 5 also includes the cointegration error of consumption, wealth and income (*cay*).

Note that because there is little evidence for the predictability of dividend growth and risk-free rate, these results can also be interpreted as followed: low housing supply elasticity (or high land value share) predicts low stock returns and high housing returns. In addition, note that low housing supply elasticity implies high housing price, and given that the intratemporal elasticity of substitution is less than one (an assumption we maintain in this paper), the nonhousing consumption share is low if the housing supply elasticity is low. This suggests that the nonhousing consumption share should *positively* predict stock returns but *negatively* predict housing returns. Interestingly, empirical results in Kwan *et al.*

(2015) indeed suggest that the coefficients on the nonhousing consumption share in predicting stock returns and housing returns have signs that are opposite[24].

2.4.4. Discussion on Stationary and Spurious Regressions

As shown in Table 2.3, the land value share is highly persistent and main contain a secular trend. However, because the dependent variables in the empirical analyses are not trending, the above results should not be viewed as spurious. The stationary of regression residuals shown in Table 2.4 may also suggest that the effects of land value share on excess returns are not spurious. On the other hand, the presence of persistence could raise the concern that the predictability results are biased in small samples. In what follows, I should argue that the biases in the results above are small.

According to Stambaugh (1999), the bias in estimating the coefficient in univariate regressions $E(\hat{\beta}-\beta)$ is: $b \times E(\hat{\phi}-\phi)$, where b is the regression coefficient of return innovations on innovations to the predictor variable[25]. For the excess stock return and the land value share, we estimate b to be equal to 0.4907, -0.1662, -1.2412, -4.1681, and -7.0592 at one-, four-, eight-, twelve-, and sixteen-

[24] Using data from Hong Kong, Kwan *et al.* (2015) compares the capacities of eight consumption-based models in accounting for the asset markets. Specifically, they regress stock returns and housing returns on predictors implied by the asset pricing theories that differ in the specifications of consumption utility and then compare the prediction errors. Their results in Table 6 and Table 7 clear show that, for any consumption-based model with housing, the coefficients on nonhousing consumption share are positive for stock return predictions but negative for housing return predictions.

[25] To be precise, Stambaugh (1999) shows that if the return regression is $r_{t+1}=\alpha+\beta x_t+u_{t+1}$ and the AR(1) process for the predictor is $x_{t+1}=\kappa+\phi x_t+v_{t+1}$, the small sample bias is $E(\hat{\beta}-\beta)$ is: $b \times E(\hat{\phi}-\phi)$, where $b=cov(u_t,v_t)/var(v_t)$.

quarter horizon, respectively. In addition, according to Kendall (1954) and Campbell (2008), the downward bias in estimates of ϕ is about $-(1+3\phi)/T$, where T is the sample size. Give the size of the quarterly sample in the paper ($T = 152$), the downward bias in estimate of ϕ is approximately equal to -0.026. Together with the multiple b, we can see that the biases in the slope coefficient are about -0.0128, 0.0043, 0.0323, 0.1084, and 0.1835 at one-, four-, eight-, twelve-, and sixteen-quarter horizon, respectively. Compared to the coefficients shown in Figure 2.2, it is clear that these biases are small.

2.5. Conclusion

This paper examines the general equilibrium effect of the housing supply elasticity on asset risk premia in a production economy with housing. It is found that, in a model where the shock to nonhousing consumption growth and the shock to its relative quantity to housing consumption are not independent, such as a production-economy setting featuring aggregate productivity shocks in this paper, adding housing to the consumption basket introduces a substitution benefit that lowers the equity risk premium. In addition, decreasing the housing supply elasticity lowers the equity risk premium through the enhanced substitution benefit but increase the housing risk premium through the elevated consumption risk. Empirical analyses using the land value share of home value as proxy for the housing supply elasticity are supportive to these model predictions. In particular, the low housing supply elasticity predicts low excess stock returns and higher excess housing returns, especially in the long-horizon (6 – 12 quarters) return forecast.

Overall, these results suggest that (1) understanding the formation of asset risk premia in a model with housing requires attentions to both specifications of investor preferences and shocks; (2) housing supply conditions would alter the asset risk premium and hence have impact on households' decisions on asset allocation; for example, this will help us to understand heterogeneity in the households' portfolio composition in relation to geographic variations in the housing supply elasticity (Lai *et al.* (2014)); (3) the decreasing of the housing supply elasticity in general could potentially explain the declining of the equity risk premium observed in the U.S. stock market; for the U.S. economy, the equity risk premium may stay low if factors such as the increasing concentration of population in large cities and increasingly-restricted regulation on home development continue to result in inelastic housing supply in general.

Given the importance of asset risk premia in terms of affecting asset prices, decisions on consumption and asset allocation and the fact that the housing supply elasticity has not only significant geographic variations but also secular time trend, establishing the linkage between the housing supply elasticity and the asset risk premium could benefit our understanding about households' attitude toward risk-taking and their investment behaviors. This paper takes the first step in this direction. However, the current study has several limitations. For example, the current model is constructed to help us to understand the theoretical connection between elasticity of housing supply and asset risk premia, and pave the way for the empirical study. Our focus on the qualitative properties of the equilibrium connection ignores the quantitative properties, e.g., the level of risk premia. In addition, the current model did not give any indications of the volatility and the cyclicity of the business cycle and housing market cycle. It would be interesting to extend the model with growth and examine both of its qualitative and quantitative properties[26].

[26] As an attempt to construct such a model, Appendix B outlines an extended model with growth for future research.

Chapter 3. Portfolio Demand and Housing Consumption Risk Hedging: Evidence from Geographic Variations in the Housing Supply Elasticity

Abstract: Using recent waves of PSID in the US, this paper shows that households living in metropolitan areas with less elastic housing supply invest a relatively larger fraction of their financial wealth in risky assets (stocks). A model with both housing and nonhousing consumption is developed to show that the optimal holding of the risky assets could be additionally motivated by households' hedging incentives against unfavorable housing price shocks. Such motive is dependent on location and household lifecycle: it is stronger in places with less elastic housing supply and for young households on the rising path of their lifecycle housing consumption profile. These findings indicate that, besides adjusting homeownership choices, households also rely on financial asset as means of hedging against housing consumption risk.

Key words: housing supply elasticity; portfolio composition; geographic variation

JEL No.: G11, R1

3.1. Introduction

In developing the intertemporal capital asset pricing model (ICAPM) where investors have incentive to hedge against changes in future investment opportunity, Merton (1973) points out that in a model with multiple consumption goods "there would be systematic effects on the portfolio demands reflecting hedging behavior against unfavorable shifts in relative consumption goods prices (i.e., in the consumption opportunity set)". Because Merton (1973) primarily focuses on investors' hedging incentive against investment opportunity and the equilibrium relationship among asset yields, he does not explore the hedging incentive against unfavorable shifts in relative prices of consumption goods and its effect on portfolio choice. In this paper, I apply Merton's insights by developing a simple two-period-two-goods model, in which households consume both nonhousing goods and housing service, and focus particularly on households' portfolio choice, e.g., the optimal share of savings investing in stocks (risk-asset share).

I pay particular attention to housing because it is not only a major consumption good, but importantly, its price variations are regional so that we could test the model implications on the cross-sectional difference in households' portfolio compositions. It is shown that as argued by Merton (1973), in contrast to the single consumption good model, the optimal risk-asset share in the two-period-two-goods model contains an additional component reflecting households' incentive to hedge against unfavorable housing price risk. Naturally, almost all households are born in "short" position of housing service (Sinai *et al.* (2005)). In expecting the long position of housing in the future, households try to reduce housing price risk by buying correlated assets now and selling them in the future when housing is needed. This is called "consumption hedge". Similar to the investment opportunity hedging components of risk-asset share in Merton (1973) that move with the covariance between asset returns and changes in state variables, it is shown that the consumption hedging components of risk-asset share here moves with the covariance between housing price and risky stock returns. Because national economic shocks will translate mostly into price effect rather than quantity changes in areas with less elastic housing supply, the covariance between housing price and the national stock market is higher in less elastic areas. Therefore, the model predicts that households living in less elastic areas will invest relatively more in risky asset as a mechanism of self-insurance against housing price risk. In addition, because the hedging incentive also depends on housing consumption profiles, which in turn depends in household's lifecycle, the effect of housing supply elasticity on risk-asset share is more pronounced for young households, who are facing steeper housing consumption plans.

Using geographic variations in the housing supply elasticity to account for variations in the covariance, this paper provides evidence that lower housing supply elasticity encourages more investment in stock shares. With the 2011 wave of the Panel Study of Income Dynamics (PSID) in combination with regional data and Saiz (2010)'s measure of housing supply elasticities in 269 MSAs, it is found that one unit decrease in the housing supply elasticity is associated with 2.4% increase in the risk-asset share of household's financial assets. The result is robust to an alternative measure of housing supply elasticities, viz., undevelopable land shares in MSAs, which are also estimated by Saiz (2010). The undevelopable land shares are generally regarded as exogeneous to regional factors so that it overcomes the potential endogeneity problem casued by omitted or unobservable MSA characteristics. Because the undevelopable land share is negatively correlated with the housng supply elasticity, the result shows that it has significantly positive effect on households' risk shares: 10% increases in the undevelopable land share are corrrelated with 1% increases in the risk-asset share. Given the low risk-asset share on average, the marginal effects are not economically insignificant. Consistent with the theory, it is also found that the covariance between housing price growth and S&P returns has positive effect on risk-asset share. In addition, the results from subsample regressions suggest that the risk-asset shares of young households are indeed more affected by the local housing supply conditions.

As a robustness check, I also construct an unbalanced short panel using 2001, 2005, and 2009 waves of the PSID and run pool cross-section regressions with MSA $\times$ year dummies. The time variation of the housing supply elasticity in the panel overcomes the potential self-selection problem in the cross-section regressions that less risk-averse households would self-select into big MSAs which generally have low housing supply elasticities. In addition, the MSA $\times$ year dummies control for any MSA and year effects so that the endogeneity problem caused by omitted or unobservable MSA characteristics is absent in the pool regressions. However, because of the MSA $\times$ year dummies, we cannot identify pure MSA level factors such as the housing supply elasticity. Therefore, I instead focus on lifecycle implications of the effect of housing supply elasticities on the risk-asset share of young households using the panel. As expected, it is found that, due to the housing consumption hedging incentive, young households living in less elastic MSAs are induced to hold relatively higher risk-asset shares. This is consistent with results using 2011 wave of the PSID. In addition, the differences in risk-asset shares of young households in MSAs with elastic and inelastic housing supply decreases with the cut-off age that defines young households, as the theory predicts that the difference for young households who have steepest housing consumption profiles should be the largest.

The present paper first contributes to the literature by showing that the cross-sectional differences in households' portfolio compositions can be explained by the hedge incentive against housing price risk, which is dependent on location and household lifecycle. These findings echo Merton (1973)'s view on portfolio demand for consumption hedging. In addition, by connecting the households' asset allocation choice to the housing supply elasticity, this paper contributes to the growing literature that brings finance and urban economics together (e.g., Ortalo-Magne *et al.* (2010)). The findings in this paper suggest a promising demand for financial products to hedge housing consumption risk, such as futures contracts linked to regional housing price indexes (Case *et al.* (1993)). Such demand is likely to increase as population becomes more concentrated in larger and denser metropolitan areas, where the housing supply elasticity tends to be lower.

The paper proceeds as follow: Section 3.2 is a brief review on related papers. Section 3.3 presents a simple two-period-two-goods model to motivate the empirical study. Section 3.4 contains discussion about data and variable construction, empirical model and results, and robustness check. Finally, Section 3.5 concludes.

3.2. Related Literature

The paper first closely relates to the growing literature that aims to understand the households' hedging incentive against housing consumption risk and the subsequent consequences on homeownership, housing price, housing consumption, and the risk-return relationship for housing etc. For example, Sinai *et al.* (2005) show that the incentive to hedge housing price risk by owning a house makes the probability to own and the price-to-rent ratio higher in markets with more volatile housing rent; Han (2008, 2010) find that stronger hedging incentives (e.g., steeper future housing consumption plans) induce larger housing demand (e.g., size of housing), *ceteris paribus*. In addition, Han (2013) shows that the hedging incentive to own helps to explain the negative risk-return relationship for housing observed in some MSAs in US. In a spatial equilibrium setting, Ortalo-Magne *et al.* (2010) demonstrates that the hedge demand depends on the covariance between the agents' earnings and local equilibrium rents, and has consequences on location choice and investment in local real estate.

By taking particular focuses on households' asset allocation, the present paper complements the aforementioned studies by showing that households not only rely on owning more housing asset as a way to hedge against housing price risk, they are also trying to further eliminate housing price risk by investing in national stocks. They especially do so if the local housing supply is less elastic so that the housing price has higher correlation with risky stock returns, and if they are young households so that they expect larger housing consumption in the future. In fact, because of high housing prices in a place with inelastic housing supply, consumption hedging with excess housing investment is costly and would not be effective.

To some extent, this research also shares some similarities with papers studying the role of nontradable goods in an open economy in explaining the "home bias", which documents the concentration of domestic assets in the portfolios held by domestic investors despite the apparent diversification gains to be had from holding foreign asset. Several papers show that the bias arise as households try to hedge the fluctuations in their consumption of nontraded goods (Baxter *et al.* (1998), Eldor *et al.* (1988), Hnatkovska (2010), and Tesar (1993)). Although this research investigates geographic variation of household's portfolio across regions within a country, it is clear that the local housing service that play crucial role in the current study is comparable to the nontraded goods in an open economy. I show that consumption hedging incentive depends importantly on the supply elasticity of the nontraded good.

The viewpoint of allocating investment in assets of which returns are correlated with housing price as means of hedging housing risk is not new in the portfolio choice literature that considers housing. In constructing the consumption-based capital asset pricing model with housing and housing transaction cost, Flavin *et al.* (2008) has pointed out that if the covariance matrix of risky asset and housing prices is not block diagonal, risky financial assets would be used to hedge the risk associated with current and future housing price. However, the same as Merton (1973), Flavin *et al.* (2008) focus not on portfolio choice but on the Euler equation of the housing CCAPM and they assume zero covariance between housing price and stocks. Englund *et al.* (2002), Iacoviello *et al.* (2003), and Quigley (2006) examine hedging housing risk in a mean-variance framework by allowing positions in real estate stocks, and housing price derivative instruments, but because they treat housing risk purely from the investment perspective and ignores future housing consumption needs, the consumption hedge incentive is absent in their model. This paper differs from this line of research by treating housing as a consumption good, and exploring the effect of geographic variation in the covariance matrix between regional housing prices and national stock returns on household portfolio composition choice. By focusing on cross-sectional variations in asset allocation, the present study extends the growing literature that examines the lifecycle portfolio compositions with housing (e.g., Cocco (2004), Fischer *et al.* (2013), Hu (2005), and Yao *et al.* (2005)).

Finally, it is widely recognized that the housing supply elasticity varies substantially across regions that are due to either differences in either physical and geographical constraints or regulatory practices (Glaeser *et al.* (2008), Green *et al.* (2005), Ortalo-Magné *et al.* (2011), Quigley *et al.* (2005), and Saiz (2010)), and that the price elasticity of housing supply play important role in affecting housing price level, volatility, persistence of housing market cycles, and urban form (Ferreira *et al.* (2011) , Fu *et al.* (2010), Glaeser *et al.* (2006), Glaeser *et al.* (2008), Huang *et al.* (2012), and Paciorek (2013)). Because the volatility and boom-bust cycles of housing market crucially depend on the local housing supply elasticity (Ferreira *et al.* (2011), Glaeser *et al.* (2008), Huang *et al.* (2012), and Paciorek (2013)), the covariance between housing price shock and national stock market returns, and hence the households' portfolio composition, also depend on housing supply elasticity. However, little research has been undertaken to understand the implications of local housing supply elasticity for household's asset allocation choice. The present study fills the gap by establishing the link between geographic variation in the housing supply elasticity and households' asset allocation.

3.3. Conceptual Framework

This section presents a conceptual framework to motivate the empirical study in this paper. A simple two-period-two-goods model is developed to show how the housing supply elasticity affects the optimal holding of risky assets. It is shown that unlike the single consumption good model, the optimal holding of risky assets is additionally motivated by households' hedging incentives against unfavorable housing price shocks.

Let time be discrete and consider a household living for two periods[27]. Assume the household has saving S at time t measured after consumption. At time $t+1$, the household consumes all his/her wealth which include gross return on investment and his/her labor income. Therefore, in order to maximize the time $t+1$ consumption utility, the household optimally allocates the saving S to two financial securities that are available to him/her: a risky asset (stocks) with gross return R_{t+1}^{s} and a risk-free asset (Treasury bills) with constant gross return R_f. Assume the return on stocks is log-normal with mean r^s and variance σ_s^2:

$$r_{t+1}^{s} = \ln\left(R_{t+1}^{s}\right) \sim N(r^s, \sigma_s^2) \tag{27}$$

Because we are interest in households' hedging incentives against the risk of relative consumption goods prices, assume households consume both nonhousing (numeraire) good C_{t+1} and housing service H_{t+1} at time $t+1$ and have CRRA preferences:

[27] A two-period model is simple to handle, and importantly, sufficient for the purpose of this study. Fama (1970) has noted that as long as the preference and the investment opportunity sets are invariant with state and time, the intertemporal portfolio choice problem of infinite horizon or multiple periods can be treated as if the households have single period utility function. The model settings in this paper satisfy the conditions stated in Fama (1970).

$$U\left(C_{t+1}, H_{t+1}\right) = \frac{\left(\left(C_{t+1}\right)^{\omega}\left(H_{t+1}\right)^{1-\omega}\right)^{1-\gamma}}{1-\gamma} \tag{28}$$

where $0<\omega<1$ and $\gamma>0$ represent the consumption share of nonhousing goods and relative risk aversion, respectively. Without loss of generality, households are assumed to obtain housing service from rental market at price P_{t+1}^{h} (housing rent), which is also log-normal[28]:

$$p_{t+1}^{h} = \ln\left(P_{t+1}^{h}\right) \sim N(p^{h}, \sigma_{h}^{2}) \tag{29}$$

where p^{s} and σ_{h}^{2} are the mean and variance of housing rent. The housing rent p_{t+1}^{h} and the risky return r_{t+1}^{s} are not independent of each other, but have contemporaneous covariance σ_{sh}. As will be shown below, the covariance between the housing rent and the risky return is critical for understanding the regional difference in households' asset allocations.

Let α be the proportion of saving invest in stocks over stocks plus bills, and $R_{t+1}^{p} = \alpha\left(R_{t+1}^{s} - R^{f}\right) + R^{f}$ the return on investment portfolio at time $t+1$.

[28] Because housing price is the capitalization of housing rent, assuming households own housing rather than renting will not alter the theoretical results of the paper but render the model less tractable. For simplicity, we assume households in the model are housing renters. In addition, the distributional assumption of P_{t}^{h} is not critical. We assume log-normal distribution because it allows us to obtain analytical solution of optimal risk-asset share later. When analytical solutions are not practical, we will assume the price of housing service (housing rent) follows gamma distribution, as suggested by empirical evidences.

Following Deaton (1991) and Carroll (1997), denote the sum of gross return on investment R_{t+1}^{p} and labor income Y_{t+1} by *cash-on-hand* at time $t+1$: $R_{t+1}^{p}S+Y_{t+1}$. Therefore, the household's optimization problem could be summarized as:

$$\begin{aligned} &\max_{\alpha,\, C_{t+1},\, H_{t+1}} \; E_t\left(U(C_{t+1},H_{t+1})\right) \\ &s.t.:\; C_{t+1}+P_{t+1}^{h}H_{t+1}=R_{t+1}^{p}S+Y_{t+1} \end{aligned} \tag{30}$$

According to (28) and (30) , we can have the household's indirect utility function:

$$\begin{aligned} &E_t\left(V(P_{t+1}^{h},R_{t+1}^{p};\mathrm{Y}_{t+1})\right) \\ &=\left(1-\gamma\right)^{-1}\omega^{\omega(1-\gamma)}\left(1-\omega\right)^{(1-\omega)(1-\gamma)}S^{1-\gamma}E_t\left(\left(R_{t+1}^{p}+\frac{\mathrm{Y}_{t+1}}{S}\right)^{1-\gamma}\left(P_{t+1}^{h}\right)^{(\omega-1)(1-\gamma)}\right) \end{aligned} \tag{31}$$

Equation (31) clearly shows that the optimal risk-asset share α^{*} depends on the joint distribution of risky return R_{t+1}^{s}, the income to saving ratio Y_{t+1}/S, and the housing rent P_{t+1}^{h}. Because of the power of summation in the conditional expectation, we are not able to obtain analytical solution for α^{*} with the distributional assumptions given above. To obtain analytical solution for a better understanding, first ignore labor income and assume $Y_{t+1}=0$ for certain. The assumption of zero labor income will be relaxed, and α^{*} will be solved numerically to examine how the results are sensitive to assumptions about labor income risk.

3.3.1. Assume No Labor Income ($E_t[Y_{t+1}] = 0, Var_t[Y_{t+1}] = 0$)

If $Y_{t+1} = 0$ for certain, the indirect utility function can be simplified as:

$$
\begin{aligned}
&E_t\left(V(p_{t+1}^h, r_{t+1}^p; 0)\right) \\
&= (1-\gamma)^{-1} \omega^{\omega(1-\gamma)} (1-\omega)^{(1-\omega)(1-\gamma)} S^{1-\gamma} E_t\left(\exp\left((1-\gamma) r_{t+1}^p + (\omega-1)(1-\gamma) p_{t+1}^h\right)\right)
\end{aligned}
\tag{32}
$$

where r_{t+1}^p is the log of the portfolio return: $r_{t+1}^p = \ln\left(R_{t+1}^p\right)$. According to the approximation method in Campbell *et al.* (2001), the log of portfolio return r_{t+1}^p can be approximated as $r_{t+1}^p \approx \alpha(r_{t+1}^s - r^f) + r^f + \frac{1}{2}\left(\alpha - \alpha^2\right)\sigma_s^2$, where $r^f = \ln\left(R^f\right)$. Plug the approximation into (32), we can solve for the optimal risk-asset share:

$$
\alpha^* = \frac{r^s - r^f + \frac{1}{2}\sigma_s^2}{\gamma\sigma_s^2} + \frac{(1-\omega)(\gamma-1)\sigma_{sh}}{\gamma\sigma_s^2} \tag{33}
$$

Equation (33) gives us a clear understanding about the determinants of optimal risk-asset share in a model with two consumption goods. The first term in is the risk-asset share in a model without housing consumption. As expected, it increases with the risk premium $r^s - r^f + \frac{1}{2}\sigma_s^2$ but decrease with risk aversion γ and variance of stock returns σ_s^2. The second term is of our primary interest. It shows that as long as the elasticity of intertemporal substitution (EIS) is not too high

($\gamma > 1$)[29], the optimal risk-asset share increases with the covariance between the housing rent p_{t+1}^h and the risky return r_{t+1}^s:

$$\frac{\partial \alpha^*}{\partial \sigma_{sh}} = \frac{(1-\omega)(\gamma-1)}{\gamma \sigma_s^2} > 0 \tag{34}$$

A simple calibration with $\omega = 0.8$, $\gamma = 5$ and $\sigma_s^2 = 0.16^2$ suggests that the marginal effect of 10 basis points increase in σ_{sh} on the optimal risk-asset share is 0.625%. As we will discuss in Section 3.4, the mean risk-asset share of the U.S. households' portfolio is relatively low and the standard deviation of σ_{sh} during 1992-2012 across MSAs is 26 basis points, so the marginal effect is not economically unimportant.

If denote η as the housing supply elasticity, we are mainly interested in the sign of $\frac{\partial \alpha^*}{\partial \eta}$. Because demand shocks will mostly translate into price effects rather than quantity effect in less elastic areas, the covariance between housing price and stock returns should be higher in these areas, implying $\frac{\partial \sigma_{sh}}{\partial \eta} < 0$. This is consistent with

[29] The existing studies on the elasticity of intertemporal substitution generally support that EIS<1, implying γ>1. For instance, Havránek *et al.* (2013) collect 2,735 estimates of the elasticity of intertemporal substitution in consumption from 169 published studies that cover 104 countries, and find the mean reported estimates of EIS is 0.5. Among the six countries that they have more than 50 estimates, Havránek *et al.* (2013) find the second largest EIS (0.6) for the US, following the largest EIS (0.9) for Japan. The mean reported estimate of EIS for the US is also close to the baseline calibration of 2/3 used by Smets *et al.* (2007).

the simulation results in Leung *et al.* (2011), who construct a multi-region general equilibrium model and show numerically that the correlation between the regional house price and the contemporary period stock price is singnificantly higher in region with higher housing stock adjustment costs. The negative correlation between housing supply elasticity and the covariance between housing price and stock returns also has empirical support, as we will see in the next section where I discuss the data. Therefore, we have:

$$\frac{\partial \alpha^*}{\partial \eta} = \frac{\partial \alpha^*}{\partial \sigma_{sh}} \frac{\partial \sigma_{sh}}{\partial \eta} = \frac{(1-\omega)(\gamma-1)}{\gamma \sigma_s^2} \frac{\partial \sigma_{sh}}{\partial \eta} < 0 \tag{35}$$

$\frac{\partial \alpha^*}{\partial \sigma_{sh}} > 0$ in (34) and $\frac{\partial \alpha^*}{\partial \eta} < 0$ in (35) provide the basis for the main tests in the paper.

Although the optimal risk-asset shares in (34) and (35) do not depend on housing consumption H_{t+1} in the simple two-period model, the negative effect of the housing supply elasticity on risk-asset share due to housing consumption hedging should be more pronounced for young households, who are more likely to trade up to a bigger home and hence have stronger hedge incentives. This lifecycle implication will be tested in the empirical analysis too.

We abstract from labor income in the above discussion. If the labor income is riskless, and is high in less elastic areas as suggested by spatial equilibrium model, households will have a higher risk-asset share because of the substitutability of labor income for bills (Bodie *et al.* (1992)). This is still the case even the labor income has idiosyncratic risk (Viceira (2001)). Therefore, the effect of housing supply elasticity on risk-asset share through the consumption hedging incentives will be confounded by the substitution effect. However, because the real labor income normalized by housing price is not necessarily higher in less elastic areas, omitting the substitution effect may not seriously bias the estimation of hedging effect. Nonetheless, household family income is controlled in the empirical models in order to take into account the possible substitution effect.

Abstracting from labor income will be more problematic if the labor income shocks contain region-specific components so that the covariance between the labor income shock and unexpected stock returns are systematically different across regions. For example, it is likely that regions may respond differently to nationwide forces, such as monetary and fiscal policies, changes in relative price of energy, and technological innovations. Difference in industrial mixes may also contribute to regional labor income cycles. As shown in Viceira (2001), whenever the return on risky asset is positively (negatively) correlated with labor income, the optimal risk-asset share contains a nonzero component representing negative (positive) hedging demand for stocks. Therefore, if the regional labor income shock has specific covariance with the national stock returns, it systematically affects households' asset allocation in that region.

However, there is little empirical evidence on whether and to what extent the labor income differs across regions, let alone how the differences are correlated with the housing supply elasticity. To examine the sensitiveness of the model prediction to the abstraction of labor income risk, we next assume risky nonzero income and experiment with various assumptions on the joint distribution of risky return R_{t+1}^{s}, the income to saving ratio Y_{t+1}/S, and the housing rent P_{t+1}^{h}.

3.3.2. Assume Risky Nonzero Labor Income ($E_t[Y_{t+1}] \neq 0, Var_t[Y_{t+1}] \neq 0$)

The housing supply elasticity η could not only affect the covariance between housing rent and stock return σ_{sh}, but also have impact on the covariance between labor income and stock return σ_{sy}, and the covariance between labor income and housing rent σ_{yh}. Without clear theoretical and empirical guidance on the calibrations of the changes of σ_{sy} and σ_{yh} with respect to η relative to changes of σ_{sh}, assume both σ_{sy} and σ_{yh} increase in pace with σ_{sh} when η deceases, or $\frac{\partial \sigma_{sy}}{\partial \eta} = \frac{\partial \sigma_{yh}}{\partial \eta} = \frac{\partial \sigma_{sh}}{\partial \eta} > 0$. This is an extreme assumption about how the joint distribution of risky return, labor income, and housing rent are affected by η.

With the above assumption, we can solve for the optimal risk-asset share numerically. As before, we set $\omega = 0.8$, $\gamma = 5$ and $\sigma_s^2 = 0.16^2$. The expected risky return r^s and return on risk-free asset are set to be $r^s = 0.08$, $r^f = 0.03$, respectively. In addition, assume the housing rent P_{t+1}^{h} and the labor income normalized by saving Y_{t+1}/S follow gamma distributions. To obtain the shape and scale parameters of these gamma distributions, I first set mean and variance of P_{t+1}^{h} and Y_{t+1}/S. Using the 2011 wave of PSID, I find the cross-sectional mean and variance of housing rent P_{t+1}^{h} and income to total wealth ratio Y_{t+1}/S are $P^{h}=39.05$, $\sigma_{P^h}^{2}=701.74$, $YS=8.36$, and $\sigma_{YS}^{2}=6608.90$, respectively[30]. These imply the shape and scale parameters of the gamma distributions being $\kappa_{P^h}=2.13$, $\theta_{P^h}=17.97$, $\kappa_{YS}=0.01$, and $\theta_{YS}=709.54$. With these parameters at hand, I then increase the Spearman's rank correlation between the housing rent and stock returns $\rho\left(R_{t+1}^{s},P_{t+1}^{h}\right)$ from 0 to 1 that can be considered as being caused by the decreases in the housing supply elasticity, and at the same time set the Spearman's rank correlation between the housing rent and the normalized labor income $\rho\left(P_{t+1}^{h},Y_{t+1}/S\right)$ and the Spearman's rank correlation between the stock return and the normalized labor income $\rho\left(R_{t+1}^{s},Y_{t+1}/S\right)$ being the same as $\rho\left(R_{t+1}^{s},P_{t+1}^{h}\right)$ so that all of them increase in the same pace[31]. For a given set of the Spearman's rank correlations, I first translate them into Pearson's correlation, and then simulate correlated multivariate normal random variables 10^6 times. These random variables

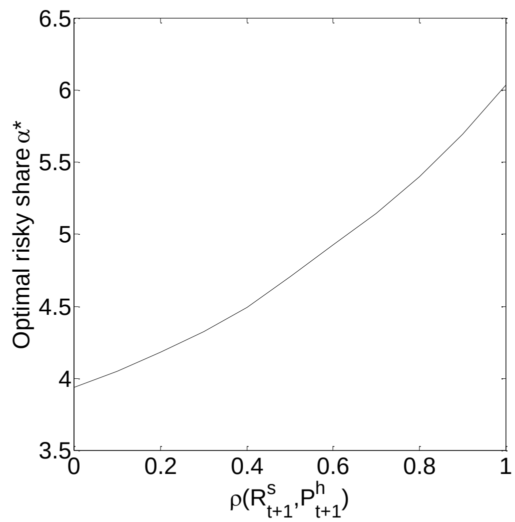
[30] The housing rent is deflated by CPI-U 2010 average (1982-84=100), and the income to total wealth ratio are restricted to samples with positive income and total wealth. Although the cross-sectional distribution may be poor estimate of the distribution of time series data, which are our interest here, the exact shape and scale parameters of the distributions are not critical in the numerical exercise because what matter are the correlations among the time series, which we set exogenously in order to examine the sensitiveness of the results to the changes in correlations.
[31] Because we fix the variance of the variables, the covariances are determined by the correlations.

are transformed to follow log normal distribution (R^s_{t+1}) and gamma distributions (P^h_{t+1} and Y_{t+1}/S) before calculating the expectation in equation (31). Finally, the optimal risk-asset share that maximizes the expectation is solved using the golden search method[32].

Figure 3.1 depicts how the optimal risk-asset share α^* changes as $\rho\left(R^s_{t+1}, P^h_{t+1}\right)$ increases from 0 to 1. Note that in this figure, $\rho\left(R^s_{t+1}, Y_{t+1}/S\right)$ and $\rho\left(P^h_{t+1}, Y_{t+1}/S\right)$ increase in the same pace as $\rho\left(R^s_{t+1}, P^h_{t+1}\right)$ increases. As can be seen, the optimal risk-asset share α^* still increases with the $\rho\left(R^s_{t+1}, P^h_{t+1}\right)$ even under the extreme assumption that $\rho\left(P^h_{t+1}, Y_{t+1}/S\right)$ and $\rho\left(R^s_{t+1}, P^h_{t+1}\right)$ increase simultaneously. This is consistent with the prediction of the simple model without labor income or with riskless labor income.

[32] Matlab code for numerically solving the optimal risk-asset share is available upon request.

Figure 3.1 Optimal risk-asset share when labor income is risky



Notes: The figure shows how the optimal risk-asset share moves with the correlation between the housing price changes and risky returns when labor income is risky. We solve for the optimal risk-asset share $\boldsymbol{\alpha}^*$ in equation for $\boldsymbol{\rho}(\boldsymbol{R}^s_{t+1}, \boldsymbol{P}^h_{t+1})$ increasing from 0 to 1, with $\boldsymbol{\rho}(\boldsymbol{P}^h_{t+1}, \boldsymbol{Y}_{t+1}/\boldsymbol{S})$ and $\boldsymbol{\rho}(\boldsymbol{R}^s_{t+1}, \boldsymbol{Y}_{t+1}/\boldsymbol{S})$ increase simultaneously. For each set of correlations, we search for α^* that maximizes the conditional expectation in equation (31), which is evaluated as the mean of 10^6 simulations. In the simulation of random variable, $\boldsymbol{R}^s_{t+1}$ follows log normal distribution $\boldsymbol{r}^s_{t+1} = \mathbf{ln}(\boldsymbol{R}^s_{t+1}) \sim \boldsymbol{N}(\mathbf{0.08}, \mathbf{0.16^2})$, while $\boldsymbol{P}^h_{t+1}$ and $\boldsymbol{Y}_{t+1}/\boldsymbol{S}$ follow gamma distribution with parameters $\kappa_{P^h} = 2.13$, $\theta_{P^h} = 17.97$, $\kappa_{YS} = 0.01$, and $\theta_{YS} = 709.54$.

Overall, the simple two-period-two-goods model in this section suggests that, for the purpose of hedging housing price risk, the positive covariance between housing price growth and risky returns induce households to have higher demand for risky assets. Because the covariance depends on local housing supply elasticity, the model predicts households living in areas with less elastic housing supply should invest a relatively larger fraction of their financial wealth in risky assets. In addition, because housing consumption demand depends on lifecycle, the effect of housing supply elasticity on risk-asset share should be more pronounced for young households. We proceed to empirically test the model predictions in the next section[33].

3.4. Empirical Evidence

3.4.1. Data and Variable Construction

To test the implications of the model, various sources of data on both household and MSA levels are used. Household level risk-asset shares are extracted from recent waves of the Panel Study of Income Dynamics (PSID)[34]. They are the dependent variable in the empirical analyses. To test equation (35), I use the housing supply elasticities in MSAs estimated by Saiz (2010) as explanatory variable[35]. Saiz (2010) also provides the undevelopable land shares in MSAs. Because of its exogeneity, the undevelopable land shares have advantage of being less likely correlated with omitted regional factors. Therefore, I also use them as a proxy for housing supply

[33] The discussions above about demand for risky asset for hedging housing price risk seem to be most relevant to positive housing price shocks. Some may concern that negative shocks and the kinked function of housing supply that is elastic w.r.t. positive shocks but inelastic w.r.t. downward shocks could invalidate the predictions. Admittedly, the hedging incentive may be absent if households expect negative housing price shocks so that the housing supply elasticity would not have impact on portfolio choice. However, positive shocks are dominant in the housing market so the housing supply elasticity takes effect. In addition, since it can be shown that the volatility of housing price is proportional to demand volatility even when housing development is irreversible Guthrie (2010), the kinked housing supply does not necessarily mean asymmetric housing price risk. Therefore, as long as upside risk dominates, the model predictions still hold with kinked housing supply.

[34] PSID data are public available at http://psidonline.isr.umich.edu/. However, "PSID-Geocode Match Files" that identify the location (e.g., MSAs) where respondents live are restricted. Some of the data used in this analysis are derived from Restricted Data Files of the Panel Study of Income Dynamics, obtained under special contractual arrangements designed to protect the anonymity of respondents. These data are not available from the authors. Persons interested in obtaining PSID Restricted Data Files should contact through the Internet at PSIDHelp@isr.umich.edu.

[35] I thank Albert Saiz for sharing his data.

elasticities. In addition, I construct the covariance between housing price growth and S&P returns in MSAs, and test directly its effect on risk-asset shares in accordance to equation (34). The definition of variables and the data sources are discussed in details below.

Risk-asset share and other household characteristics: Starting at the household level, the risk-asset share and other household characteristics are obtained from the PSID. The PSID contains rich household level information about asset holdings and many other household characteristics including age, gender, education, family income, and so on. Following the common practice in empirical literature, I define risk free savings as the sum of cash, checking and savings, bond and insurance, and refer the risky assets as the sum of holdings of stocks and mutual funds. The risk-asset share is extracted from the PSID as the ratio of risky assets to financial assets, which is the sum of the risky assets and risk free savings.

Although the PSID is a longitudinal study that tracks households and their descendants over time, I focus on the most recent wave of the PSID for the year of 2011 in the empirical analysis. This is not only because our primary focuses are on the cross-sectional variation of households' risk-asset shares, but also because we lack reliable time-series measure of housing supply elasticities. While Saiz (2010)'s measure of housing supply elasticities are widely used in the literature (see e.g., Huang *et al.* (2012), Mian *et al.* (2011) , and Paciorek (2013)), they are cross-sectional. As a result, I choose the 2011 wave of PSID because the year of survey is closer to the year in which Saiz (2010) estimate the MSA housing supply elasticities.

As a robustness check of the results from the 2011 wave of PSID, I also use the 2001, 2005 and 2009 waves of PSID to construct alternative samples of unbalanced short panel and run pool cross-section regressions [36]. The data structure of unbalanced short panel has the advantage of allowing us to add MSA $\times$ year dummies in the empirical model. These dummies help to control for unobservable MSA characteristics and year effects at the expense of not identifying pure MSA level factors such as housing supply elasticity. Therefore, I will focus the lifecycle implications of the model in the pool cross-section regressions by comparing young households' portfolio composition across MSAs. More about the robustness check with alternative waves of the PSID will be discussed in Section 3.4.3.

Housing supply elasticity and undevelopable land share: At the MSA level, we rely on the housing supply elasticities in MSAs estimated by Saiz (2010). Using the satellite-generated data, Saiz (2010) computes the undevelopable land shares in MSAs, which are the percentage of land within a 50-km radius from the metropolitan central cities that are unsuitable for housing development because of water bodies, wetlands, and steep slopes. Saiz (2010) then provides estimates of housing supply elasticity measure as functions of both physical and regulatory

[36] Again, we do not use all the consecutive biennial waves of PSID from 2001 to 2009 in constructing the panel because we are more interested in cross-sectional variation. In addition, because there is little time-variation in the housing supply elasticity within short periods, using all consecutive waves of the PSID adds little value for testing cross-sectional differences.

constraints for 269 major metropolitan areas in the U.S. It ranges from 0.60 to 12.15, with smaller value indicating lower housing supply elasticity.

Because the housing supply elasticities are considered to be endogenous to population growth and other MSA economic factors, the empirical results can be biased if there are omitted and unobservable MSA characteristics that are correlated with both of housing supply elasticity and households' investment behavior. In contrast, the undevelopable land share, as an argument of the housing supply elasticity function, is generally regarded as purely exogenous to most regional economic factors. The exogeneity of the undevelopable land share is useful for dealing with the potential endogeneity problem caused by omitted and unobservable MSA characteristics. Therefore, the undevelopable land shares are also used as a proxy for housing supply elasticities in the cross-section regressions. Clearly, the undevelopable land shares in MSAs are negatively correlated with the housing supply elasticity, as argued by Saiz (2010).

Covariance between housing price growth and risky returns: I construct the covariances between housing price growth in MSAs and the S&P 500 returns so that we can examine their correlation with the housing supply elasticities and undevelopable land shares, as well as directly test their effect on households' risk-asset shares. As shown in the conceptual framework, it is the negative correlation with the covariance between housing price growth and risky returns through which the housing supply elasticity affects risk-asset share. We want to see whether the negative correlation exists in the data. To construct the covariances, the quarterly Federal Housing Finance Agency Purchase-Only Indexes in 100 largest MSAs spanning 1991Q1-2014Q2 (hereafter the FHFA indexes) are used. I define the yearly housing price growth in MSAs as the log difference of the 4th quarter FHFA indexes. Data on S&P composition price and dividend from Rober Shiller's website are used to measure stock returns. For each MSA, I use the yearly housing price growth and stock returns from 1992 to 2013 to calculate the covariance between housing price growth and S&P 500 returns.

Other MSA economic variables: To control for other regional factors, I obtain local unemployment rates by MSA from Bureau of Labor Statistics, population by MSA from U.S. Census Bureau, and GDP by MSA from Bureau of Economic Analysis. I will use the average unemployment rate, the average population growth, and the average GDP growth in MSAs from 2007 to 2011 as controls for regional factors. Using the 2011 PSID, I also calculate mean of risk free savings to wealth ratio for households whose family income is below the 5% percentile of the empirical family income distribution in each MSA, and use it to control for the potential regional difference in the social security network.

As discussed in the conceptual framework section, regional-specific income risks may correlate with housing supply elasticities and affect households' portfolio choice. However, it is difficult to find appropriate measures of income risk in empirical analysis[37]. In the cross-section regressions, I follow Carlino *et al.* (2001) and use the standard deviation of real per capita income from 2008 to 2012 by MSA as control for potential regional difference in labor income risk[38]. In the robustness check with alternative samples, the MSA $\times$ year dummies in the pool cross-section regressions are able to control for MSA characteristics including labor income risks.

With the aforementioned datasets at hand, I then merge the MSA level data with the PSID using location identifier in PSID-Geocode Match Files. Because the risk-asset share is not defined for households with zero financial assets, I restrict the sample to households with positive financial asset. I also transform some household control variables to minimize the effect of outliers. For example, I divide household yearly income and total wealth by the sample mean in order to reduce their magnitude, and I take the log of household mortgage payment after adding a small number (0.01). In addition, I multiply the covariance by 10,000 so that its unit is basis point. Other household controls such as education, gender, marital status, occupation, race, health status, and family composition change are dummy or category variables and are left as they are. After steps of merging and cleaning the data, the final usable

[37] Most studies about the effect of labor income risk on portfolio choice are theoretical and rely on numerical simulation (see e.g., Bodie *et al.* (1992), Viceira (2001), and Polkovnichenko (2007)). Some empirical analyses rely on rely on self-reported indicators of income risk (e.g., Guiso *et al.* (1996)).

[38] Although nominal per capita income by MSA is available since 1969, the real per capita income by MSA that is adjusted by regional price parities (RPP) is only recently available from 2008 to 2012 at www.bea.gov/regional/index.htm

cross-sectional sample based on 2011 wave of PSID consists of 4,799 households from 285 MSAs[39].

Table 3.1 presents summary statistics of the combined data. In the table, I define MSAs with inelastic housing supply as those with housing supply elasticities lower than 25% percentile of the empirical distribution, and otherwise elastic MSAs. As can be see, the mean of risk-asset share in MSAs with inelastic housing supply is 0.20. It is relative higher than the mean of risk-asset share in elastic MSAs, which is 0.17. We can also observe from Table 3.1 that the undevelopable land share and the covariance between housing price growth and S&P stock returns are negatively correlated with the housing supply elasticity: In MSAs with inelastic housing supply, the means of the undevelopable land share and the covariance are both much higher. This relation is more evident in Table 3.2, which presents the correlation matrix of these three variables. It suggests that the price growths indeed have higher covariance with the national stock returns in less elstic MSAs, consistent with the implications of the two-region general equilibrium model in Leung *et al.* (2011). Importantly, this lends support to the conjecture in the conceptual framework because the negative correlation between the housing supply elasticity and the covariance between housing price growth and risky returns are critical for the effect of housing supply elasticity on the hedging demand for stocks to operate.

[39] However, the exact sample size will vary with the specifications of empirical model.

Table 3.1 Summary statistics

	All samples			Samples in elastic MSAs			Samples in inelastic MSAs		
	N	Mean	Std Dev	N	Mean	Std Dev	N	Mean	Std Dev
Panel A: Summary statistics of household variables (2011 PSID)									
risk-asset share	4,799	0.18	0.31	3,159	0.17	0.3	1,640	0.2	0.32
age	4,799	46.21	16.56	3,159	45.71	16.43	1,640	47.17	16.79
occupation	4,781	9.08	6.16	3,150	9.04	6.15	1,631	9.17	6.18
health status	4,781	2.39	1	3,148	2.39	1	1,633	2.38	1.01
college degree	4,799	0.43	0.5	3,159	0.43	0.5	1,640	0.44	0.5
family size	4,799	2.57	1.41	3,159	2.61	1.43	1,640	2.48	1.38
gender (1=male, 2=female)	4,799	1.27	0.45	3,159	1.27	0.44	1,640	1.28	0.45
married	4,799	0.61	0.49	3,159	0.62	0.49	1,640	0.6	0.49
race	4,764	1.51	1.1	3,132	1.47	1.02	1,632	1.58	1.24
mortgage payment (log)	4,755	4.41	4.75	3,126	4.52	4.74	1,629	4.18	4.77
income	4,799	1.3	1.55	3,159	1.3	1.64	1,640	1.31	1.35
total wealth	4,799	1.48	5.59	3,159	1.51	6.05	1,640	1.42	4.56
family composition change	4,799	0.88	1.71	3,159	0.93	1.76	1,640	0.79	1.58
Panel B: Summary statistics of MSA variables									
housing supply elasticity (Saiz (2010))	205	2.5	1.45	154	2.96	1.38	51	1.09	0.27
undevelopable land share (Saiz (2010))	205	0.26	0.21	154	0.17	0.13	51	0.53	0.17
cov(housing price growth, S&P returns)	60	26.98	25.99	41	20.2	21.41	19	41.61	29.4
mean risk free savings of low income households	256	0	6.26	209	0.41	0.8	47	-1.82	14.49
average unemployment rate (2007-2011)	258	6.65	1.78	207	6.57	1.89	51	6.96	1.21
average population growth (2007-2011)	282	0.01	0.02	232	0.01	0.02	50	0.01	0.01
average GDP growth (2007-2011)	271	0.51	1.85	221	0.67	1.78	50	-0.2	1.97
s.d. of real per capita income (2008-2012)	271	0.11	0.06	221	0.1	0.06	50	0.11	0.06

Notes: This table shows the summary statistics of the combined data using 2011 wave of PSID and MSA level datasets. For the 2011 PSID, the sample is restricted to those with positive financial asset. The housing supply elasticities and the undevelopable land shares in MSAs are obtained from Saiz (2010). MSAs are defined as elastic (inelastic) if housing supply elasticities are below (above) the 25% percentile of the empirical distribution. Occupation, health status, college degree, male, married, race are all dummy or category variables with respect to household head. A small number (0.01) is added to mortgage payment before taking log. Income and total wealth are relative measures in the sense that they have been divided by sample means. FHFA Purchase-Only Indexes are used for obtaining housing price growth in MSAs. Covariances between housing price growth and S&P returns are calculated using time series from 1992 to 2012.

Table 3.2 Correlation matrix

	housing supply elasticity	undevelopable land share	cov(housing price growth, S&P returns)
housing supply elasticity (Saiz (2010))	1.000 [247]		
undevelopable land share (Saiz (2010))	-0.619(0.00) [247]	1.000 [247]	
cov(housing price growth, S&P returns)	-0.359(0.01) [53]	0.359(0.01) [53]	1.000 [77]

Notes: The table shows the pairwise correlations between variables at the MSA level. The housing supply elasticity and undevelopable land share are obtained from Saiz (2010). FHFA Purchase-Only Indexes are used for obtaining housing price growth in MSAs. Covariances between housing price growth and S&P returns are calculated using time series from 1992 to 2012. Figures in brackets are number of observations, and figures in parentheses are significant levels.

3.4.2. Empirical Methodology and Results

This section presents the empirical specifications and results of testing the cross-sectional difference in households' portfolio composition caused by households' hedging demand for stocks, which in turn depends on housing supply elasticities and household lifecycles. I first examine the model implications on cross-sectional difference in risk-asset share, and then proceed to subsample regressions by age group in order to investigate the lifecycle implications.

In accordance to equation (35) and equation (34), the baseline specification are cross-section regressions with the risk-asset share of a household i in MSA k $\left(\alpha_{i,k}\right)$ as dependent variable and the MSA level variables of interest (x_k) as explanatory variable. Specifically, the linear model of the following form is estimated:

$$\alpha_{i,k} = \beta_0 + \beta_1 x_k + \Theta \mathrm{X}_i + \Phi Z_k + \varepsilon_{i,k} \tag{36}$$

where x_k is either the housing supply elasticity, the undevelopable land share, or the covariance between housing price growth and S&P returns. X_i is a vector of household level controls including household head's age, family size, log of mortgage payment, income relative to sample mean, total wealth relative to sample mean, dummy or category variables for household head's education, gender, marital status, occupation, race, health status, and family composition change. Z_k is a vector of MSA level controls including average unemployment rate, average population growth, and average GDP growth from 2007 to 2011, the standard deviation of real per capita income from 2008 to 2012, and the mean risk free saving to financial asset ratio of low income households.

The conceptual framework in Section 3.3 predicts that households live in less elastic MSAs should invest relative more in stocks, indicating a negative sign on β_1 when the explanatory variable is the housing supply elasticity. Because the undevelopable land share is negative correlated with the housing supply elasticity, the sign on β_1 should be positive when x_k is the undevelopable land share. Finally, according to equation (34), the sign on the covariance between housing price growth and S&P returns should be positive.

Table 3.3 reports the estimation results, with column from (1) to (3) representing results from using either housing supply elasticities, undevelopable land shares, or the covariance between housing price growth and S&P returns as explanatory variable, respectively. Column (1) of Table 3.3 shows that, as expected, the sign on the housing supply elasticity is significantly negative at 5% level: One unit decrease in the housing supply elasticity is associated with about 2.4% increases in the risk-asset share. Given the mean risk-asset share across households in the sample is 18%, the marginal effect of the housing supply elasticity on the risk-asset share is economically substantial.

To address the potential endogeneity problem caused by omitted and unobservable MSA characteristics that are correlated with both the housing supply elasticity and households' investment behavior, I replace the housing supply elasticity with the undevelopable land share and repeats the analysis. Because the undevelopable land share is negatively correlated with the housing supply elasticity, and other mechanisms affecting the risk-asset share are unlikely to be systematically different in MSAs with high and low undevelopable land share, the measurement of land constraint is a valid proxy of the housing supply elasticity and should have positive effect on risk-asset shares. Column (2) of Table 3.3 reports the regression results of using the undevelopable land share as explanatory variable. Consistent with the theory, the sign on the undevelopable land share is significantly positive. The coefficient suggests that an increase of 10 percentage point of the undevelopable land share will result in 1% increases in the risk-asset share. This lends further support to model predictions.

Table 3.3 Main results in cross-section regressions (PSID 2011)

	Dependent variable: risk-asset shares		
	(1)	(2)	(3)
β_1: housing supply elasticity (Saiz (2010))	-0.024**		
	(0.0099)		
β_1: undevelopable land share (Saiz (2010))		0.10**	
		(0.048)	
β_1: cov(housing price growth, S&P returns)			0.0014***
			(0.00051)
head's age	0.0029***	0.0029***	0.0014
	(0.00074)	(0.00074)	(0.00087)
head has college degree	0.078***	0.078***	0.069***
	(0.013)	(0.014)	(0.017)
family size	-0.017***	-0.018***	-0.013
	(0.0053)	(0.0054)	(0.0091)
head is female	-0.031*	-0.033*	-0.036*
	(0.018)	(0.018)	(0.022)
married	0.024	0.023	0.014
	(0.025)	(0.025)	(0.032)
mortgage payment	0.0036**	0.0038**	0.0019
	(0.0016)	(0.0016)	(0.0022)
income	0.025**	0.025**	0.042***
	(0.012)	(0.012)	(0.014)
total wealth	0.0082	0.0082	0.014***
	(0.0052)	(0.0052)	(0.0049)
mean risk free savings of low income households	-0.00035	-0.00039	-0.00068**
	(0.00036)	(0.00037)	(0.00033)
average unemployment rate (2007-2011)	-0.0046	-0.0031	-0.014**
	(0.0072)	(0.0068)	(0.0068)
average population growth (2007-2011)	0.18	0.28	1.15
	(0.53)	(0.54)	(0.79)
average GDP growth (2007-2011)	0.0027	0.0057	0.013
	(0.0047)	(0.0051)	(0.0096)
s.d. of real per capita income (2008-2012)	0.048	0.013	-0.035
	(0.18)	(0.18)	(0.21)
Constant	-0.058	-0.18*	0.033
	(0.11)	(0.10)	(0.078)
Other individual controls	Yes	Yes	Yes
State dummy	Yes	Yes	Yes
N	3397	3397	1695
Adj. R-Square	0.20	0.20	0.26

Notes: This table reports the regression results of equation (36) using 2011 wave of PSID in combination with MSA level data. Estimations are by OLS. Column (1) to (2) uses the housing supply elasticity and undevelopable land share in Saiz (2010) as explanatory variables; column (3) uses the covariance between housing price growth and S&P returns as explanatory variable. Other individual controls include dummies for household head's occupation, health status, family composition change, and race. Standard errors are clustered at MSAs and reported in parentheses. *, **, and *** denote statistically significant at the 10%, 5% and 1% level, respectively.

The final column of Table 3.3 report the estimation results using the covariance between housing price growth and S&P return as explanatory variable. This is a direct test of model implications expressed in equation (34). Consistent with the model prediction, the covariance has significantly positive effect on risk-asset share. 10 basis point increases in the covariance between housing price growth and S&P returns correspond to 1.4% increases in risk-asset share of households' portfolio, a number not too far away from model prediction in Section 3.1[40].

The effects of other control variables on risk-asset share are generally in line with existing findings. For example, the risk-asset shares increase with age, education, family income, and total wealth, and decrease with family size and if household head is female. Interestingly, as have been found in Heaton *et al.* (2000) and Cocco (2004), higher mortgage payment is positively correlated with higher risk-asset share, suggesting households may finance stock investment via mortgage debt. Other regional variables generally have expected sign but are not statistically significant. For instance, the sign on mean risk free saving to financial asset ratio of low income households is negative, suggesting the worse the social security the lower the risk-asset share; high unemployment rate is associated with low risk-asset share, while high population and GDP growth are related to high risk-asset shares. However, although the coefficients are not significant, the regional labor income risk seems to have positive effect on risk-asset share.

[40] Recall that the simple calibration in Section 3.3.1 suggests 10 basis point increases in the covariance will cause risk-asset share increases 0.625%.

Possible explanation could be measurements error in labor regional labor income risks, or because high regional labor income risks are correlated with high labor incomes that are not captured by the included regional factors (e.g., GDP growth).

The channel through which the housing supply elasticity has impact on the risk-asset share depends on not only the negative effect of the housing supply elasticity on the covariance between housing price growth and S&P returns, but also the housing consumption plan of the households. If households currently own housing and have litter demand for upsizing housing consumption, their hedging demands for stocks are weak. Therefore, the effect of housing supply elasticity on risk-asset share, if any, should be moderate for old households, who are likely to downsizing their housing consumption in the future. Actually, the effect of the housing supply elasticity on risk-asset share could be opposite to the model prediction for old households because they are expecting a "short" position of housing in the future[41]. In contrast, young households have steepest lifecycle housing consumption profile, so they should have strongest hedging demand for stocks in expecting upcoming increases in housing consumptions. The dependence of hedging motive on the households' lifecycle housing consumption

[41] Empirical evidence on the downsizing of housing at the old age is mixed. Chiuri *et al.* (2010) use 60 microeconomic surveys on about 300,000 individuals residing in 15 OECD countries to explore the pattern of elderly homeownership and find that ownership rates decline considerably after age 60 in all countries. Using U.S. data, Venti *et al.* (2004) find that old households are unlikely to discontinue homeownership and liquidate home equity to support general nonhousing consumption needs. Fang (2009) show that the empirical lifecycle housing consumption profile is not hump-shape for U.S. household. Rather, it first increases monotonically and then flattens out at age about 60.

plan suggests the effect of housing supplies on risk-asset shares should be more pronounced for young households.

To test the lifecycle implications of the model, I divide the samples into two different age groups: young households and old households. According to Fang (2009)'s estimates, the consumption profile for housing of the U.S. households starts to flattens out at about age 60 (see Figure 3 and Figure 4 therein). Therefore, I define young households as those of which the head's age is not greater than 60, and repeat analyses of equation (36) using the age subsamples. Table 3.4 shows the regression results, with the first three columns for young household samples and the last three columns for old household samples. Again, I use either the housing supply elasticity, the undevelopable land share, or the covariance between housing price growth and S&P returns as explanatory variable. The estimates for these variables have expected sign and are all significant for young household subsamples. They are all insignificant for the old age group.

Table 3.4 Results in cross-section regressions by age subsamples (2011 PSID)

	Dependent variable: risk-asset shares					
	subsample: households' age<61			subsample: households' age>=61		
	(1)	(2)	(3)	(4)	(5)	(6)
β_1: housing supply elasticity (Saiz (2010))	-0.021**			-0.041		
	(0.0098)			(0.026)		
β_1: undevelopable land share (Saiz (2010))		0.15***			-0.044	
		(0.050)			(0.11)	
β_1: cov(housing price growth, S&P returns)			0.0012**			0.0017
			(0.00048)			(0.0013)
N	2735	2735	1393	662	662	302
Adj. R-Square	0.20	0.20	0.23	0.18	0.18	0.34

Notes: This table reports the regression results of equation (36) by age subsamples using 2011 wave of PSID in combination with MSA level data. Estimations are by OLS. Controls are the same as those in Table 3.3. To conserve space, only selected coefficients are reported. Standard errors are clustered at MSAs and reported in parentheses. *, **, and *** denote statistically significant at the 10%, 5% and 1% level, respectively.

3.4.3. Robustness Check with Alternative Waves of the PSID

I use earlier waves of the PSID to provide a robustness check of results from 2011 wave of the PSID. Specifically, I use the 2001, 2005 and 2009 waves of the PSID to construct an unbalanced short panel, and then run pool cross-section regressions using this alternative samples. Due to lack of time-series measure of housing supply elasticity, we combine information on MSA population and undevelopable land shares to create a time-varying dummy variable ($inelasticMSA_{k,t}$) to indicate whether a MSA has inelastic housing supply: the dummy equals to one if both the undevelopable land share and the population in the MSA is above the 75% percentile of their empirical distributions, and zero otherwise. As discusses in Saiz (2010), the undevelopable land share is more likely to play role in affecting housing supply elasticity in MSAs with large population. Although it is basically time-invariant, the populations in MSAs would have significant relative changes in the last decade[42]. Therefore, the inelasticity dummy variable $inelasticMSA_{k,t}$ should be able to at least partially capture variations in the housing supply elasticity across MSAs and over time.

Due to data availability, we do not collect regional economic factors in the pool cross-section regressions. Instead, I run pool cross-section regression with MSA × year dummies. This allows us to control for MSA and year effects at the expense of not identifying pure MSA level factors such as housing supply elasticity. However, we can explore the lifecycle implications of housing supply elasticity on risk-asset shares, following the identification strategy in Sinai *et al.* (2005). Explicitly, the econometric model is specified as follow:

$$\alpha_{i,k,t} = \beta_0 + \beta_1 Young_{i,t} + \beta_2 Young_{i,t} \times inelasticMSA_{k,t} + \Theta X_{i,t} + \Phi MSA_k \times Year_t + \varepsilon_{i,k,t} \tag{37}$$

where is $\alpha_{i,k,t}$ the risk-asset share of a household living in MSA k at time t. The $Young_{i,t}$ a dummy variable that equal one if the household head's age is blow a

[42] For instance, 2010 census special reports on "Patterns of Metropolitan and Micropolitan Population Change" suggests that there are substantial geographic variation in population growth in the U.S. between 2000 and 2010, with rapid growth in some areas of the country and sizable declines in others. According to the report, the fastest-growing metro areas were located in either the South or the West, with fastest population gainers led by Palm Coast, FL, and followed by St. George, UT; Las Vegas-Paradise, NV; Raleigh-Cary, NC; and Cape Coral-Fort Myers, FL. Two metro areas in the southern states of Louisiana and Arkansas (New Orleans and Pine Bluff, respectively) and three areas located partially or entirely in the states of Ohio, Pennsylvania, and West Virginia (Youngstown-Warren-Boardman, OH-PA; Johnstown, PA; and Steubenville-Weirton, OH-WV) are the fastest-declining metro areas. The report is available at https://www.census.gov/prod/cen2010/reports/c2010sr-01.pdf.

cut-off age that defines young and zero if the household head's age is above 40[43]. The $inelasticMSA_{k,t}$, as discussed above, indicates whether a MSA k at time t belongs to the inelastic MSA group. It equals $landshareHigh_k \times populationHigh_{k,t}$, with $landshareHigh_k$ or $populationHigh_{k,t}$ being indicator that equals one if the corresponding variable in the MSA (undevelopable land share or the population) is above the 75% percentile of its empirical distribution. X_i includes the same household level controls as before, and $MSA_k \times Year_t$ is a set of MSA × year dummies.

The specification (37) compares the portfolio composition of young households and old households. Because the base group is the old households ($Young_{i,t} = 0$), β_1 tells the difference in risk-asset shares between these two demographic groups, *ceteris paribus*. More importantly, the specification (37) also compares the difference in risk-asset shares of young households across MSAs. The coefficient β_2 on the interaction term suggests additional difference in risk-asset shares of young households if they are from MSAs with inelastic housing supply. A positive β_2 suggests they hold a higher risk-asset share in comparison with those in elastic MSAs.

[43] Note that we will experiment with different cut-off ages in order to examining how β_2 is sensitive to the cut-off age defining young households. For the purpose of comparability, we exclude observations with household heads' age falling in the range between the cut-off age and 40, but keep households with age above 40 as the base group.

Because Fang (2009) shows that for both renter and home owners, the housing consumption plans have steepest slop before the age of 30. To more sharply compare the difference in risk-asset shares of young households due to housing consumption hedging demand, I first set the cut-off age at 31 and report estimation results in Table 3.5. The significant and negative β_1 shows that young households hold a smaller fraction of their financial portfolio in risky stocks than old households. However, the significant and positive β_2 suggests that the gap is narrower if these young households live in MSAs with inelastic housing supply. All else equal, young households from MSAs with inelastic housing supply have risk-asset share about 4.4% higher than their counterparts from MSAs with elastic housing supply (about a quarter of the sample mean share of risky asset in household financial portfolio). The result provides significant evidence on the importance of consumption risk hedging motive of financial investment choices.

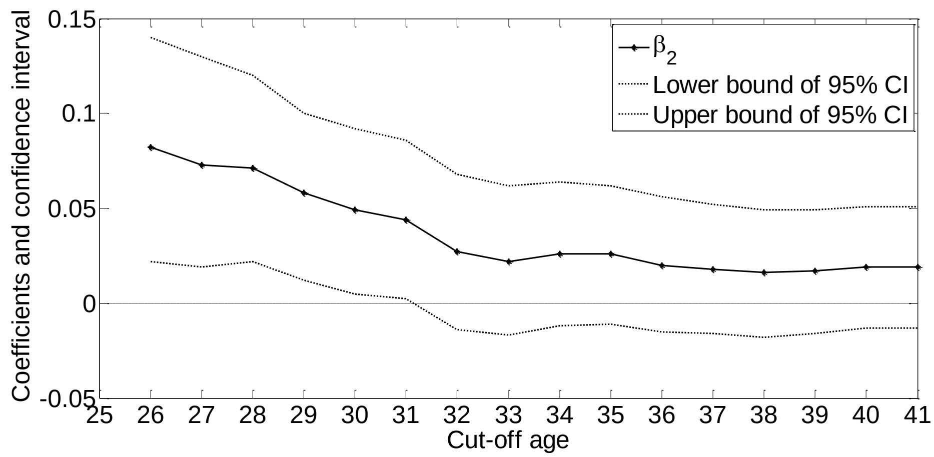
Table 3.5 Results with alternative waves of the PSID (PSID 2001, 2005, 2009)

	Dependent variable: risk-asset shares
	(1)
β_1: Young households (age<31)	-0.078***
	(0.017)
β_2: Young households (age<31) × inelastic MSA	0.044**
	(0.021)
Age	0.0021***
	(0.00048)
Head has college degree	0.13***
	(0.0094)
Family size	-0.0099***
	(0.0034)
Head is female	-0.0080
	(0.014)
Married	0.042***
	(0.013)
Mortgage payment	0.0034***
	(0.00099)
Income	0.0080***
	(0.0029)
Total wealth	0.0040**
	(0.0017)
Other individual controls	Yes
MSA × year dummies	Yes
N	10692
Adj. R-Square	0.23

Notes: This table reports the regression results of equation (37) using the alternative samples of unbalanced short panel constructed with the 2001, 2005 and 2009 waves of PSID. Estimation is by OLS. The base group consists of old households whose heads have age greater than 40. Other individual controls include dummies for household head's occupation, health status, family composition change, and race. Standard errors are clustered at MSA × year cells and reported in parentheses. *, **, and *** denote statistically significant at the 10%, 5% and 1% level, respectively.

Figure 3.2 depicts how the coefficient β_2 changes if we increase the cut-off age from 26 to 41. Consistent with the theory, β_2 decreases with the cut-off age. For the youngest households (households' age < 26), because they have steepest housing consumption plan and the incentive to hedging housing consumption risk, the relative difference in risk-asset shares in MSAs with elastic and inelastic housing supply are largest and significantly different from zero. As the cut-off age increases, both the magnitude and significance level of the difference decrease.

Figure 3.2 $\boldsymbol{\beta_2}$ and confidence intervals in pool cross-section regressions



Notes: The figure shows the sensitiveness of the coefficient on the interaction term of young household indicator and inelastic MSA indicator, β_2 in equation (37), to the cut-off age defining young households. The coefficients β_2 measure the extent to which young households in MSAs with inelastic housing supply have a greater share of risky asset in investment portfolio for housing consumption hedge. A lower bound of 95% confidence interval above zero indicates that the difference is significant at 5% level. MSAs are defined as inelastic if both the undevelopable land share and population are above the 75% percentile of their empirical distributions. Standard errors are clustered at MSA × year cells and are used to construct the 95% confidence intervals.

3.5. Conclusion

Households on a rising housing consumption path face price risks in places where housing supply is inelastic. In a simple two-period-two-good model we show that holding risky assets, whose returns are correlated with housing price shocks, provides consumption hedging benefits. In particular, this consumption benefit is greater in places where housing supply elasticity is lower. Using recent waves of PSID, this paper shows that households living in MSAs with inelastic housing supply indeed hold relatively higher fraction of their financial portfolio in stocks. Consistent with the household lifecycle consumption theory, it is further shown that the effect of housing supply elasticity on risk-asset share is more pronounced for young households on the rising path of their housing consumption profile.

The present study contributes to a growing literature on household housing consumption hedging behavior. In addition to its effects on homeownership (Sinai *et al.* (2005)), timing and size of housing consumption (Han (2008, 2010)), and the price-rent ratio in the housing market (Sinai *et al.* (2005) and Han (2013)), the housing consumption hedging incentive also significantly affect households' investment choices, especially for young households. These findings suggest that financial innovations, such as housing futures and option contracts based on regional home price indices tied to regional housing price index (Case *et al.* (1993)), has a promising demand. As population continues to concentrate in larger and denser metropolitan areas, where housing supply elasticity tends to be lower, such demand is likely to increase. However, improving the attractiveness of a financial product to households is always challenging and would be a promising area for future research.

Chapter 4. Competitive Consumption Spending and Labor Supply: Evidence from Regional Differences in Sex Ratio in China

Abstract: Wei *et al.* (2011) show that the substantial increase in household saving in China since late 1990s may have to do with a rising male-female sex ratio. They find a higher sex ratio in a region in China makes parents with a young son save more for the son's expenses, such as wedding and education, and also invest in a bigger home, to help the son compete in local marriage market. The present study investigates the impact of sex ratio on young males' consumption spending and labor supply behavior. It is hypothesized that marriage market competition makes young males spend more where the sex ratio is higher – they may do so with financial support from their parents. Furthermore, young males in high sex ratio regions would also work harder, so that their earning would rise faster, in order to pay back their parents in the future. A large dataset of credit card account information of individuals across 31 provinces in China is employed to test these hypotheses. It is found that an additional percentage point in regional sex ratio of age 20 to 34 in 2005 is associated with two to three percent higher credit card balance for males in this cohort but not for females. It is also found that young males' age profile of income to be steeper in provinces with higher sex ratio. These findings are consistent with the proposed hypotheses and suggest that the rising sex ratio in China may also have contributed to China's high GDP growth through competitive consumer spending and labor supply by young males.

Key words: sex ratio imbalance, competitive spending, labor supply

JEL No.: D1; R1

4.1. Introduction

Since the introduction of the one-child policy placing birth population control in 1979, China has experience a significant rise in the sex ratio of male to female in recent decades. Especially, because of the parental preferences for sons that root in the traditional Chinese cultural norms and the adoption of Ultrasound B for prenatal screening, the sex ratio at birth has been became increasingly imbalanced since 1980s and the maleness of young adults has progressively increased. According to Edlund *et al.* (2013), the sex ratio of young adults of 16 to 25 years old rose from 1.02 to 1.06 between 1988 and 2004. This implies a very competitive marriage market for males. Even the Chinese government has recently relaxed the one-child policy, because the sex ratio imbalance at birth has been increasing steadily in the last 30 years, the sex ratio imbalance for young adults and the competitiveness of marriage market would continue to rise for at least two decades.

Motived by recent studies about the effects of rising sex ratio imbalance in China on social and economic variables such as saving rates and crime rates (Edlund *et al.* (2013); Wei *et al.* (2011)), this paper investigates males' credit card spending behavior and working efforts in associated with sex ratios. Different from Wei *et al.* (2011) which shows that parents with a son have competitive saving motive to improve their son' relative standing in the marriage market, this paper hypothesizes that the male children themselves (young adults) have competitive consumption motive in order to attract marriage partners. This hypothesis is developed based on two observations: First, the parents' competitive savings shown in Wei *et al.* (2011)

are related to their children's future consumption (especially housing consumption). it can be considered that, with financial support from their parents, young males can increase current consumption as a signal of attractiveness. Second, although it is limited by the line of credit, the use of credit cards is convenient. More importantly, because the credit market is still in the embryonic stage in China and consequently credit cards are normally marketed to individuals with relative stable and high income, the use of credit cards can signal the social-economic status of card holders (Worthington (2003), Estelami *et al.* (2007) and Worthington *et al.* (2011)). Therefore, it can be conjectured that due to the rising sex ratios that stiffen the competition in marriage market, males have competitive spending motive, and this motive induce males in places with more imbalanced sex ratios to increase the usage of credit cards and thus have higher balance in their credit card accounts. In addition, in order to pay back to their parents, e.g., via financial support for their parents' retirement, young males in high sex ratio regions would also work harder, so that we also conjecture their earning would rise faster.

Using a unique dataset of credit card account information covering 31 provinces and the inferred sex ratios for young adults in 2005 at province level based on the 2010 population census in China, this paper finds empirical evidence for a positive relationship between sex ratios and males' credit card balance and working efforts. In particular, it is shown that one percentage increase in the local sex ratio for age 20 to 34 is associated with 2% - 3% increase in males' credit card balance and about 0.44% increase in males' income path. Interestingly, the falsification exercises show that these effects are not significant or even become negative if the samples are restricted to female observations. Recognizing that the sex ratios of young adults in 2005 inferred from the 2010 population census are potentially measured with errors due to population migration and different mortality rates for males and females, an alternative measurement of the sex ratios based on both 2000 and 2010 population censuses is constructed. It is found that the results are robust to this alternative construction of sex ratios.

These findings are consistent with the view that higher sex ratios have increased females' bargaining power in the marriage market and men tend to exert every effort to increase their attractiveness to mates in response to rising pressure in the marriage market. On one hand, it is well known that an increase in the sex ratio may increase the intensity of intra-sexual competition for status to attract mates. On one hand, because a credit card applicant in China has to convince the credit card issuers that he/she can meet the payment timely with a stable income, the use of credit card serves well as signals of social status because of its visibility. Therefore, as shown by results in this paper, men present their attractiveness and social-economic status to a marriage partner through swiping credit cards. This is similar to display conspicuous consumption to signal desirable mate qualities (Griskevicius *et al.* (2007) and Sundie *et al.* (2011)). Results in this paper paint a full picture which suggests that high sex ratios increase males' demand for showy spending – i.e., spending through the use of credit cards – to improve their marriage prospect and lead males to work harder to have steeper income path. By examining the competitive spending motive of the males themselves and their effort in improving income prospects, this paper complements to Wei *et al.* (2011) and contributes to the literature by showing evidence of competitive spending motives and increased working effort induced by imbalance sex ratios. These results suggest that the high GDP growth in China may have to do with the rising sex ratio imbalance through its effect on competitive consumption spending and labor supply of young males.

The paper is organized as followed: Section 4.2 is a brief review on related works, followed by empirical analyses in Section 4.3, where data, descriptive statistics and results are discussed; as a robustness check, Section 4.4 contains a repeated empirical analyses with an alternative construction of sex ratios and discuss the results; finally, Section 4.5 concludes.

4.2. Literature Review

A first strand of literature relevant to the current study is the theory of marriage and how changes in the sex ratio affect male's and female's bargaining power that in turn have consequences on social and economic variables. Becker's model of marriage and family formation (Becker (1973); Becker (1974)) provide a theoretical framework to understand these effects. For instance, Becker notes that an increase in the sex ratio (of male to female) will increases the demand for wives as it enhances the role of female in the output of the joint production of a marriage. An important implication of Becker's theory relevant to the current study is that men are motived to invest in characteristics (e.g., education) in order to become attractive in markets where the marginal product of women in marriage is high. This implies a positive correlation between sex ratios and males' characteristics such as education and earnings. Consistent with theories that female bargaining power in the marriage market increases with sex ratios, Angrist (2002) finds higher sex ratios have a positive effect on the likelihood of female marriage and male's earnings but negative effect on female labor force participation. Similarly, Chiappori *et al.* (2002) provide both structural model and empirical evidence (using PSID data for the 1989 interview year) that rise in the sex ratios leads to a reduction in female labor supply and an increase in male labor supply.

However, while the existing theories of marriage generally focus on the bargaining position of man and woman and the intra-household allocation of resources, few studies, both theoretically or empirically, examine effects of imbalanced sex ratios on consumption. If women are scarce, the intensity of intra-sexual competition for mates will reasonably trigger men to increase their showy consumption as a signal of their attractiveness. Sundie *et al.* (2011) investigate this hypothesis and find conspicuous consumption indeed serves as a social signal directed at potential mates, although they also find men who conspicuously consume more likely target at short-term mating. On the other hand, psychologists find that a scarcity of women leads people to expect that men spend more money during courtship (such as by paying bills or engagement rings); it also leads men to discount the future more heavily and incur debt for immediate expenditures (Griskevicius *et al.* (2012)). This study uses the balance on credit card accounts as instrument for consumption and test whether it increases with sex ratios[44].

Another strand of literature works on how status competition can alter agents' consumption and saving behavior (Cole *et al.* (1992); Hopkins *et al.* (2004); Hopkins *et al.* (2006)). For instance, Cole *et al.* (1992) introduce a matching process of man and woman into the classical growth model that provides man an incentive to save beyond that captured by standard models. As pointed out by Wei *et al.* (2011), however, this strand of literature normally does not feature sex ratio imbalance. This paper empirically address how the status competition for attracting marriage partners resulted from imbalanced sex ratios affect males' spending behavior.

Finally, the rising sex ratio imbalance in recent decades in China has been attracting attention from economist and blamed for the high household savings (Wei *et al.* (2011)) and the increase in crime (Edlund *et al.* (2013)). With the examination of the effect of imbalanced sex ratios on male's credit card spending and income path, this study continues in this line of research.

4.3. Empirical Analysis

4.3.1. Data

The main source of data is a unique dataset provided by one of the major Chinese banks about individual's credit card information covering 31 provinces in China. The

[44] The credit card balance potentially captures both the men's "showy" purpose to attract mates and their desire for "immediate" gratification, which we cannot disentangle.

dataset contains individual particulars such as gender, age, marital status, education, job industry, job seniority, and the zip code of their residence. It also contains information about credit card balance, personal income, credit card line, quasi-credit card line, and loan line at the time of retrieving information, which concentrates in 2005 and 2006[45]. Initially, the dataset has about 1.5 million observations. However, we have to exclude observations having missing values in variables of interest, e.g., education, occupation, job seniority, zip code, and personal income. In addition, because the effect of imbalanced sex ratio on credit card spending, if exists, may limit to relative young people, the sample is further restricted to populations with age between 18 and 49. To reduce noise, the top 1% and the bottom 5% of samples in terms of their personal income, credit line and the top 1% of samples in terms of credit card balance are removed. These steps left us with about 50 thousands of observations in the regression analysis.

The second set of data is about regional variables that are used to control for regional effects on individual spending behavior. These variables mostly measure the economic climate in provinces, including GDP growth from 2000 to 2005, per capita income, share of labor force enrolled in social security, and share of SOE employment in total labor force in 2005 at provincial level. They are obtained from Wei *et al.* (2011), where they are used as controls for regional effects on households' saving behavior. I also obtain the average residential housing price in 2006 from

[45] Quasi-credit card is a debit card that has credit function but may require collateral (e.g., a security deposit or the guarantee of the monthly salary payment) to function as credit card. It is a transitional product at the early stage of developing credit market in China.

Chinese Real Estate Statistics Yearbook as a control for the effect of housing pricing on credit card spending[46].

The last key regional variable is the sex ratio of male to female at provincial level. Because the credit card balances are retrieved mainly in 2005 and 2006, we ideally would like to know sex ratios for each age cohort in every region during that period. However, such data are not obtainable because the Chinese population census is carried out only about every 10 years (e.g., in 1982, 1990, 2000, and 2010). In addition, the public available population census data assembly only reports the population for age cohort of every five years, e.g., 0, 1-4, 5-9, 10-14, etc. Therefore, the inferred sex ratios of age group of 20-34 in 2005 based on the sex ratios of the 25-39 cohort in 2010 population census are used in the empirical analyses[47]. As mentioned in Wei *et al.* (2011), this method of inferring the sex ratio is likely to underestimate the actual sex ratio in 2005 because the mortality rates for boys and young men are generally slightly higher than those for girls and young women. However, because the underestimations are common across all provinces, the measurement errors will only affect the constants in the regression results unless the differences in the mortality rates between young men and young women have significant provincial variation, which we do not notice any supportive anecdotes or studies.

[46] See the average residential housing price in 2006 at provincial level at: http://tongji.cnki.net/kns55/addvalue/areaindusdevelop.aspx?sicode=Z013&areacode=xj05

[47] Initially, I inferred the sex ratio of male to female for two age groups, e.g., age 20-34 and age 35-49, in 2005 from the 2010 population census by calculating the sex ratio of age groups of 25-39 and 40-54 in 2010. As it is found that sex ratios of the two age groups are higher correlated, with correlation coefficient being 0.8396, the sex ratio of the younger cohort is chosen in the empirical analyses. I have also inferred the sex ratio of a broader age group of 15-34 in 2005 from 2010 population census and found its correlation with the sex ratio of age group of 20-34 in 2005 is 0.995.

4.3.2. Descriptive Statistics

Table 4.1 describes the summary statistics of key variables. Panel A of Table 4.1 shows that the sex ratio have substantial variation across provinces. The mean of sex ratios in provinces with high sex ratio is 1.10 with standard deviation of 5%, while the mean of sex ratios in provinces with low sex ratio is 1.01 with standard deviation of 3%[48]. Examining other regional variables, we observe that provinces with high sex ratios are normally more developed regions, as evident by the mean of per capita income, GDP growth, and residential housing price. Panel B of Table 4.1 shows that the average credit card balance is 685.14 RMB in the sample. With an average annual income of 51 thousands RMB, this means that the credit card balance is about 16.1% of monthly income. Comparing the mean of credit card balances in provinces with high sex ratios and provinces with low sex ratios, we can see that the mean of credit card balance in regions with high sex ratios is clearly higher than that in regions with low sex ratios (765.91 versus 580.35). We can also see the mean of income and total line (the sum of credit card line, quasi-credit card line, and loan line) are higher in regions with more imbalanced sex ratios. This further shows that, as suggested by the summary statistics of the regional variables, provinces with high sex ratios are generally richer. These summary statistics are suggestive that people in regions with more imbalanced sex ratio have higher income and spend more using credit cards. However, the standard deviations of the credit card balances and income within each group of provinces easily overwhelm the differences between the groups of regions.

[48] The sex ratio is defined as high if it is greater than the mean of sex ratio across provinces.

Table 4.1 Summary statistics

	All samples			Samples in provinces with low sex ratio			Samples in provinces with high sex ratio		
	N	Mean	Std Dev	N	Mean	Std Dev	N	Mean	Std Dev
Panel A: Summary statistics of regional variables									
Sex ratio	31	1.05	0.06	17	1.01	0.03	14	1.10	0.05
Per capita income (in thousand)	31	5.71	3.23	17	4.69	1.27	14	6.95	4.37
GDP growth from 2000 to 2005	31	42.51	6.84	17	42.21	6.55	14	42.87	7.41
Share of labor force enrolled in social security	31	29.68	20.06	17	25.13	13.70	14	35.22	25.25
Share of SOE employment in total labor force	31	11.30	4.79	17	9.70	3.68	14	13.24	5.39
Residential housing price in 2006 (in thousand)	31	2.70	1.51	17	2.16	0.60	14	3.36	2.00
Panel B: Summary statistics of individual variables									
Credit card balance (in thousand)	937,793	685.14	2167.43	408,207	580.35	1917.05	529,586	765.91	2339.03
Income (in thousand)	279,099	50.99	52.25	121,059	39.39	44.31	158,040	59.88	56
Total line (in thousand)	1,008,223	85.24	167.55	447,451	60.8	123.83	560,772	104.74	193.34
Male	1,021,794	0.59	0.49	450,725	0.6	0.49	571,069	0.58	0.49
Single	894,878	0.27	0.44	426,695	0.21	0.41	468,183	0.32	0.47
Age	1,024,082	35.28	6.92	452,063	35.61	6.42	572,019	35.01	7.28
Education	1,024,082	0.18	0.38	452,063	0.17	0.38	572,019	0.18	0.39
Job industry	1,024,082	0.23	0.57	452,063	0.21	0.54	572,019	0.25	0.59
Job seniority	1,024,082	4.47	21.9	452,063	3.08	14.34	572,019	5.57	26.33

Note: The sample is restricted to those who are between 18 and 49 years old. Sex ratios are the ratio of male to female between the ages of 20-34 at provincial level in 2005, inferred from the 2010 population census. It is defined as high if it is higher than the mean of sex ratios across provinces. Per capita income, Share of labor force enrolled in social security, Share of SOE employment in total labor force are percentage points and data in 2005. Together with GDP growth from 2000 to 2005, these variables are obtained from Wei *et al.* (2011). Credit card balance, income, and total line (credit card line, quasi-credit card line, and loan line) are all in logs. Education, job industry, and job seniority are categorical variables, while male and single are dummies.

4.3.3. The Effect of Sex Ratio on Credit Card Balance

To test the hypothesis of competitive spending, the econometric model is specificed as followed:

$$bal_{i,j} = \beta_0 + \beta_1\ sexRatio_j + \sum \phi \Phi_i + \sum \delta \Delta_j + e_i \tag{38}$$

where $bal_{i,j}$ is the log of credit card balance of individual i in region j [49]; $sexRatio_j$ is the sex ratio in region j; β_1 is the coefficient of interest, measuring the marginal effect of $sexRatio_j$ on $bal_{i,j}$; Φ_i is a set of individual controls that include gender, marital status, education, job industry, job seniority, age, age squared, log of income, and square of log of income etc; Δ_j is a vector of regional controls that consist of GDP growth in the last 5 years, per capita income, share of labor force enrolled in social security, share of SOE employment in total labor force in 2005, and residential housing price in 2006. These regional controls are added following Wei *et al.* (2011), because they potentially affect individual's consumption and saving behavior. The equation (38) is implemented with three different set of samples. I first run regression using full samples, and then restrict to female subsamples and male subsamples. In each set of regressions, two different variables are used as control of personal income due to concerns about measurement errors in the income variable. Explicitly, Model A utilizes the log of income variable itself and Model B use the log of total line (the sum of credit card line, quasi-credit card

[49] Note that because many observations have zero credit card balance at the time of retrieving information, we add 1 to the credit card balance before taking log.

line, and loan line) as control for personal income, respectively. Since the line of credit is normally approved based on banks' discreet judgment on the prospect and riskiness of individual's income, the total line would be a good measure of the card holders' income profile.

Table 4.2 reports the regression results. It shows that for the full sample, the marginal effect of sex ratio on credit card balance is positively significant in both Model A and Model B, implying that one percentage increase in sex ratio is associated with about 1.5 percentage higher in credit card balance. Interestingly, as can be seen from the middle panel of Table 4.2, the marginal effect disappears if we restrict observations to female subsamples. In addition, the coefficients on sex ratio in Model B for female sample are negative, which suggests female in regions with more imbalanced sex ratio may have lower credit card balance. By contrast, the coefficients on sex ratio become greater and positively significant when we restrict observations to male subsamples. For model A, the marginal effect of sex ratio on male's credit card balance is now about 2.5, implying that one percentage increase in sex ratio is associated with 2.5 percentage higher in male's credit card balance. Given that the mean of male's credit card balance is 762 RMB and the mean of sex ratios across province being 1.05, a 10 percentage increase in sex ratio from 1.05 to 1.15 is at lease associated with about 190 RMB increase in monthly credit card spending. With a multiplier of the size of male population, the aggregate effect of increase in credit card spending on economy is huge.

Table 4.2 Sex ratio effect on credit card balance

	Dependent variable: log of credit card balance					
	All population		Female		Male	
	Model A	Model B	Model A	Model B	Model A	Model B
Sex ratio	1.56**	1.49***	0.19	-0.13	2.36***	2.52***
	(0.63)	(0.52)	(1.03)	(0.82)	(0.79)	(0.68)
Income (log)	0.42***		0.39***		0.44***	
	(0.020)		(0.036)		(0.025)	
Total line (log)		0.068***		0.061***		0.071***
		(0.013)		(0.020)		(0.016)
Male	0.12***	0.16***				
	(0.028)	(0.026)				
Married	0.26***	0.26***	0.36***	0.39***	0.19***	0.17***
	(0.033)	(0.032)	(0.056)	(0.053)	(0.042)	(0.040)
Age	0.014	-0.070***	-0.018	-0.078**	0.014	-0.085***
	(0.024)	(0.023)	(0.041)	(0.038)	(0.030)	(0.029)
Age squared	-0.0011***	0.000086	-0.00041	0.00043	-0.0012***	0.00015
	(0.00032)	(0.00030)	(0.00056)	(0.00051)	(0.00040)	(0.00038)
Constant	-2.16**	2.44***	0.77	4.42***	-3.21***	1.76**
	(0.90)	(0.70)	(1.54)	(1.16)	(1.12)	(0.89)
Education, Job industry, Job seniority, and Regional controls	Yes	Yes	Yes	Yes	Yes	Yes
N	65949	75511	21237	25014	44712	50497
Adj. R-Square	0.07	0.07	0.07	0.07	0.07	0.07

Note: The sample is restricted to those who are between 18 and 49 years old. Sex ratios are the ratio of male to female between the ages of 20-34 at provincial level in 2005, inferred from the 2010 population census. Model A and Model B use log of income and log of total line (credit card line, quasi-credit card line, and loan line) to control for income, respectively. Education, job industry and job seniority are categorical variables, while married is dummy variable. Regional control includes per capita income in 2005, GDP growth from 2000 to 2005, share of labor force enrolled in social security in 2005, share of SOE employment in total labor force in 2005, which are obtained from Wei *et al.* (2011), and residential housing price in 2006. Robust standard errors in parentheses. *, **, and *** denote statistically significant at the 10%, 5% and 1% level, respectively.

To further examine whether the effects of sex ratio on male's credit card spending depend on marital status, housing status and age, I implement econometric model similar to equation (38) again on male subsample but add interaction terms into the model. Explicitly, I first interact sex ratio with marital status to see whether married male is less affected by the sex ratio; I then interact sex ratio with housing status to examine whether housing owners are differently affected by the sex ratio; and finally we interact sex ratio with an age group dummy which equals one if the sample's age is between 35 and 49 and zero otherwise. The regression results of each interaction specification are reported in the left panel, middle panel, and right panel of Table 4.3, respectively. First, compared to the coefficient on the sex ratio in the right panel of Table 4.2, the coefficients on sex ratio in Table 4.3 are generally greater. This suggests that the base group, which is single male in each panel of Table 4.3, is most sensitively affected by the imbalanced sex ratio. Second, expect for the last column of Table 4.3, coefficients on the interaction term of the sex ratio with marital status, housing status, and age group are not significant, showing that the effect of sex ration on credit card spending for married males, male housing owners and older males are not statistically different from its effect on single males. Although the coefficients on the interaction terms are mostly not significant, they have expected signs. For example, married male and older are less affected by sex ratio, while male housing owners seem to react more to a rise in the sex ratio.

Overall, the results lend supports to the hypothesis that because of imbalanced sex ratios, males increase their credit card spending to attract females and improve their competitiveness in the marriage market. However, although the robust standard errors reported in Table 4.2 and Table 4.3 take into account issues concerning heteroskedasticity and lack of normality, there may be correlations among observations in one region that should not be overlooked. Table 4.4 reports results with standard errors clustered at city level. It shows that standard errors on sex ratio are larger than those reported in Table 4.2 and Table 4.3,

Table 4.3 Sex ratio effect on male's credit card balance

	Dependent variables: log of male's credit card balance					
	Sex ratio interacts with marital status		Sex ratio interacts with housing status		Sex ratio interacts with age group	
	Model A	Model B	Model A	Model B	Model A	Model B
Sex ratio	3.01***	3.22***	2.09***	2.25***	2.54***	2.96***
	(1.00)	(0.90)	(0.81)	(0.70)	(0.95)	(0.84)
Married * Sex ratio	-0.94	-0.97				
	(0.85)	(0.80)				
Own housing * Sex ratio			1.33	1.28		
			(0.99)	(0.93)		
Older (35~49) * Sex ratio					-0.30	-0.73
					(0.81)	(0.76)
Income (log)	0.44***		0.44***		0.44***	
	(0.025)		(0.025)		(0.025)	
Total line (log)		0.071***		0.071***		0.071***
		(0.016)		(0.016)		(0.016)
Age	0.013	-0.086***	0.014	-0.085***	0.049	-0.073**
	(0.030)	(0.029)	(0.030)	(0.029)	(0.034)	(0.032)
Age squared	-0.0012***	0.00017	-0.0012***	0.00016	-0.0015***	0.000039
	(0.00040)	(0.00038)	(0.00040)	(0.00038)	(0.00043)	(0.00041)
Married	0.81	0.86	0.19***	0.17***	0.19***	0.17***
	(0.90)	(0.84)	(0.042)	(0.040)	(0.042)	(0.040)
Own housing			-1.46	-1.40		
			(1.05)	(0.98)		
Older (35~49)					0.17	0.71
					(0.86)	(0.79)
Constant	-3.68***	1.23	-2.91**	2.07**	-4.12***	1.06
	(1.27)	(1.07)	(1.13)	(0.91)	(1.29)	(1.07)
Education, Job industry, Job seniority, and Regional controls	Yes	Yes	Yes	Yes	Yes	Yes
N	44712	50497	44712	50497	44712	50497
Adj. R-Square	0.07	0.07	0.07	0.07	0.07	0.07

Note: The sample is restricted to male who are between 18 and 49 years old. Sex ratios are the ratio of male to female between the ages of 20-34 at provincial level in 2005, inferred from the 2010 population census. Model A and Model B use log of income and log of total line (credit card line, quasi-credit card line, and loan line) to control for income, respectively. Education, job industry and job seniority are categorical variables, while married is dummy variable. The base group is single male (left panel), single male housing renter (middle panel), and single young male (right panel). Regional controls include per capita income in 2005, GDP growth from 2000 to 2005, share of labor force enrolled in social security in 2005, share of SOE employment in total labor force in 2005, which are obtained from Wei *et al.* (2011), and residential housing price in 2006. Robust standard errors in parentheses. *, **, and *** denote statistically significant at the 10%, 5% and 1% level, respectively.

indicating positive correlations among observations at the cluster level[50]. Unsurprisingly, most of the coefficients on sex ratio become insignificant because of cluster error given that we only have sex ratio at provincial level. Nonetheless, it is worth mentioning that for the single males, the effect of sex ratio on credit card balance remains significant at 10% level (see the first two columns of the bottom panel in Table 4.4). This suggests that marriage market competition could induce competitive spending motive, especially for single males.

Table 4.4 Regression results with cluster error

	Dependent variable: log of credit card balance					
	All population		Female		Male	
	Model A	Model B	Model A	Model B	Model A	Model B
Sex ratio	1.56	1.49	0.19	-0.13	2.36	2.52
	(1.57)	(1.70)	(1.57)	(1.39)	(1.85)	(2.06)
	Dependent variables: log of male's credit card balance (Male subsample)					
	Sex ratio interacts with marital status		Sex ratio interacts with housing status		Sex ratio interacts with age group	
	Model A	Model B	Model A	Model B	Model A	Model B
Sex ratio	3.01*	3.22*	2.09	2.25	2.54	2.96
	(1.55)	(1.90)	(1.87)	(2.13)	(1.81)	(2.05)

Notes: This table reports the standard errors of the coefficients on sex ratio corrected for clustering at city level. The upper panel corresponds to Table 4.2 and the bottom panel corresponds to Table 4.3. Other controls are the same as the corresponding tables but omitted. Clustered standard errors in parentheses. *, **, and *** denote statistically significant at the 10%, 5% and 1% level, respectively.

In the next section, we proceed to test the second hypothesis that males in regions with more imbalanced sex ratio also have to working harder to repay the debt and sustain the standard of livings.

[50] The standard errors in Table 4.4 are clustered at the first three digits of zip code. According to the coding rule of zip code in China, this can be interpreted as clustering at city level.

4.3.4. The Effect of Sex Ratio on Income Path

The econometric model of testing the effect of sex ratios on income path is specified as followed:

$$\begin{aligned} inc_{i,j} = \beta_0 + \alpha_0 \ Age_i + \beta_1 \ Age{*}D_j^{hsr} + \sum \phi \Phi_i \\ + \sum \alpha_j province_j + Age_i * RegionalControls + e_{i,j} \end{aligned} \tag{39}$$

where $inc_{i,j}$ is the log of income, measured by income variable itself and total line in Model A and Model B, respectively; D_j^{hsr} is a dummy variable that equals one if the sex ratio in region j is greater than the mean of sex ratios across provinces and zero otherwise; the coefficient β_1 measures the slope of income path on top of α_0 if the observation is from provinces with more imbalanced sex ratio. Similar to equation (38), I include other individual controls such as marital status, education, job industry, job seniority in Φ_i. I use the dummy $province_j$ to control for regional fixed effect so α_j captures differences in the intercept. To address the concerns that the sex ratios may capture the effect of omitted regional variables on the income-age profile, the age variable is also interacted with regional controls, such as GDP growth, share of labor force enrolled in social security and housing price.

I run regression with female and male subsamples. We expect β_1 to be significantly positive for male, suggesting that the income path for male in provinces with more imbalanced sex ratio is steeper. The regression results of the above specification are reported in Table 4.5. For the first two columns of Table 4.5, we can see that the coefficients on “Age × High sex ratio dummy” are not significant for female subsamples. The sign is even negative in Model A. In contrast, the coefficient on the interaction term in Model A is significantly positive for the male subsample. The result suggests that the slope of male’s income path in provinces with more imbalanced sex ratios is about 0.44% steeper. When the income is measured using total credit line, the coefficient on the interaction term is also not significant for male subsample, but the probability of nil effect is much lower in comparison with female subsample. These results suggest that males in regions with high sex ratio have steeper age-income profile.

Table 4.5 Sex ratio effect on income path

	Dependent variable: log of income			
	Female subsample		Male subsample	
	Model A	Model B	Model A	Model B
Age	0.0030	0.010	-0.011***	0.014**
	(0.0048)	(0.0076)	(0.0038)	(0.0054)
Age × High sex ratio dummy	-0.00047	0.0030	0.0044***	0.0034
	(0.0024)	(0.0039)	(0.0016)	(0.0024)
Constant	9.70***	9.89***	9.10***	9.90***
	(0.19)	(0.28)	(0.12)	(0.18)
Married	Yes	Yes	Yes	Yes
Education	Yes	Yes	Yes	Yes
Job industry	Yes	Yes	Yes	Yes
Job seniority	Yes	Yes	Yes	Yes
Age * GDP growth	Yes	Yes	Yes	Yes
Age * Share of labor force enrolled in social security	Yes	Yes	Yes	Yes
Age * Housing price	Yes	Yes	Yes	Yes
Regional fix effect	Yes	Yes	Yes	Yes
N	28779	33398	61888	69419
Adj. R-Square	0.26	0.23	0.23	0.20

Note: The sample is restricted to those who are between 18 and 49 years old, with first two columns for female subsample and the last two columns for male subsample. Sex ratios are the ratio of male to female between the ages of 20-34 at provincial level in 2005, inferred from the 2010 population census. The high sex ratio dummy equals one if the sex ratio is higher than the mean of sex ratios across provinces. Education, job industry and job seniority are categorical variables, while married is dummy variable. Model A and Model B use log of income and log of total line (credit card line, quasi-credit card line, and loan line) as the dependent variable, respectively. Regressions in the first two columns interact age with high sex ratio dummy only, regressions in the third and four column allow age interact with other individual controls, and finally regressions in the last two columns further interact age with regional controls. For the regional variables, GDP growth is the mean growth from 2000 to 2005 and share of labor force enrolled in social security is data in 2005, both of which are obtained from Wei *et al.* (2011); the residential housing price is data in 2006. Robust standard errors in parentheses. *, **, and *** denote statistically significant at the 10%, 5% and 1% level, respectively.

4.4. Robustness Check with Alternative Construction of Sex Ratio

In this section presents empirical results with alternative construction of sex ratio as a robustness check. In the last section, the sex ratios of age group of 20-34 in 2005 is inferred from the 2010 population census only. This method of inferring sex ratios overlooks population migrations and may result in measurement error in the inferred sex ratios. By contrast, I construct the sex ratios of age group of 20-34 in 2005 based on both the 2000 population census and the 2010 population census and use them in the empirical analyses in this section. Specifically, since people aged 20-34 in 2005 are 15-29 years old in 2000 and 25-39 years old in 2010, I sum up the population of aged 15-29 in the 2000 population census and the population of aged 25-39 in the 2010 population census in each province for male and female, respectively, and then divide the sum of male by the sum of female to arrive the sex ratio:

$$\textit{sexRatio} \text{ of age 20-34 in 2005} = \frac{\text{male age 15-29 in 2000 + male age 25-39 in 2010}}{\text{female age 15-29 in 2000 + female age 25-39 in 2010}} \quad (40)$$

Using both 2000 census and 2010 census to construct the sex ratio has advantage over inferring it from either wave of the population censuses. First, because young adults aged about 18 are very likely to leave the province of birth due to college enrollment after high school and college graduates aged between 20 and 30 are also mobile to seek employment after graduating from college, inferring sex ratio from either the 2000 census or the 2010 census overlooks the potential bias caused by migrations. Second, as mentioned in Wei *et al.* (2011), due to different mortality rates for males and females, inferring the sex ratio from either the 2000 census or the 2010 census is likely to underestimate or overestimate the actual sex ratio in 2005. Therefore, because the credit card information in the sample is archived in 2005 and 2006, which are in the middle of 2000-2010, constructing the sex ratio of aged 15-34 in 2005 using both waves of population census can average out measurement errors caused by migrations or different mortality rates for males and females.

The analyses of model specification (38) for credit card balance and model specification (39) for income path using the alternative construction of sex ratios and report the empirical results are reported in Table 4.6 and Table 4.7, respectively[51]. It can be seen that results shown in Table 4.6 are similar to these in Table 4.2. Importantly, Table 4.6 shows that only males' credit card balances are significantly affected by the sex ratio and the coefficients on the sex ratio for the female subsamples is negative (although they are not significant). Similar to results in Table 4.5, the coefficients on "Age × High sex ratio dummy" in Table 4.7 are positively significant (except the last column) for male subsample when income is measured by

[51] In repeating model specification (38), we have also further interacted sex ratio with dummies for marital status, housing status and age group and find the results are similar to those shown in Table 4.3. To conserve space, the results with interaction terms are not reported.

the income variable itself, suggesting that the slope of male's income path in provinces with more imbalanced sex ratios is steeper.

Table 4.6 Sex ratio effect on credit card balance: alternative construction of sex ratio

	Dependent variable: log of credit card balance					
	All population		Female		Male	
	Model A	Model B	Model A	Model B	Model A	Model B
Sex ratio	0.97	1.60**	-1.47	-0.81	2.45**	3.17***
	(0.96)	(0.80)	(1.58)	(1.24)	(1.21)	(1.05)
Income (log)	0.42***		0.39***		0.44***	
	(0.020)		(0.036)		(0.025)	
Total line (log)		0.068***		0.060***		0.072***
		(0.013)		(0.020)		(0.016)
Male	0.13***	0.16***				
	(0.028)	(0.026)				
Married	0.25***	0.26***	0.35***	0.39***	0.19***	0.17***
	(0.033)	(0.032)	(0.056)	(0.053)	(0.042)	(0.040)
Age	0.014	-0.070***	-0.018	-0.078**	0.015	-0.084***
	(0.024)	(0.023)	(0.041)	(0.038)	(0.030)	(0.029)
Age squared	-0.0011***	0.000082	-0.00041	0.00043	-0.0012***	0.00014
	(0.00032)	(0.00030)	(0.00056)	(0.00051)	(0.00040)	(0.00038)
Constant	-1.61	2.35***	2.31	5.06***	-3.28**	1.18
	(1.12)	(0.90)	(1.89)	(1.45)	(1.41)	(1.16)
Education, Job industry, Job seniority, and Regional controls	Yes	Yes	Yes	Yes	Yes	Yes
N	65949	75511	21237	25014	44712	50497
Adj. R-Square	0.07	0.07	0.07	0.07	0.07	0.07

Note: The sample is restricted to those who are between 18 and 49 years old. Sex ratios are the ratio of male to female between the ages of 20-34 at provincial level in 2005, inferred from both the 2000 population census and the 2010 population census (see equation (40)). Model A and Model B use log of income and log of total line (credit card line, quasi-credit card line, and loan line) to control for income, respectively. Education, job industry and job seniority are categorical variables, while married is dummy variable. Regional control includes per capita income in 2005, GDP growth from 2000 to 2005, share of labor force enrolled in social security in 2005, share of SOE employment in total labor force in 2005, which are obtained from Wei *et al.* (2011), and residential housing price in 2006. Robust standard errors in parentheses. *, **, and *** denote statistically significant at the 10%, 5% and 1% level, respectively.

Table 4.7 Sex ratio effect on income path: alternative construction of sex ratio

	Dependent variable: log of income			
	Female subsample		Male subsample	
	Model A	Model B	Model A	Model B
Age	0.0036	0.010	-0.012***	0.012**
	(0.0048)	(0.0075)	(0.0037)	(0.0053)
Age × High sex ratio dummy	0.00040	0.0028	0.0038***	0.00099
	(0.0021)	(0.0036)	(0.0015)	(0.0022)
Constant	9.67***	9.90***	9.12***	9.98***
	(0.18)	(0.27)	(0.12)	(0.18)
Married	Yes	Yes	Yes	Yes
Education	Yes	Yes	Yes	Yes
Job industry	Yes	Yes	Yes	Yes
Job seniority	Yes	Yes	Yes	Yes
Age * GDP growth	Yes	Yes	Yes	Yes
Age * Share of labor force enrolled in social security	Yes	Yes	Yes	Yes
Age * Housing price	Yes	Yes	Yes	Yes
Regional fix effect	Yes	Yes	Yes	Yes
N	28779	33398	61888	69419
Adj. R-Square	0.26	0.23	0.23	0.20

Note: The sample is restricted to those who are between 18 and 49 years old, with first two columns for female subsample and the last two columns for male subsample. Sex ratios are the ratio of male to female between the ages of 20-34 at provincial level in 2005, inferred from the 2010 population census. The high sex ratio dummy equals one if the sex ratio is higher than the mean of sex ratios across provinces. Education, job industry and job seniority are categorical variables, while married is dummy variable. Model A and Model B use log of income and log of total line (credit card line, quasi-credit card line, and loan line) as the dependent variable, respectively. For the regional variables, GDP growth is the mean growth from 2000 to 2005 and share of labor force enrolled in social security is data in 2005, both of which are obtained from Wei *et al.* (2011); the residential housing price is data in 2006. Robust standard errors in parentheses. *, **, and *** denote statistically significant at the 10%, 5% and 1% level, respectively.

4.5. Conclusion

This paper empirically investigates whether males will increase their credit card spending and work harder to earn higher income in response to an increase in the sex ratios (of male to female). Empirical analyses using a unique dataset of credit card account information covering 31 provinces and the regional variations in the sex ratio of young adults in China provide empirical evidence for a positive relationship between sex ratios and males' credit card balance and working efforts. Instead of the competitive saving motive of parents with son in Wei *et al.* (2011), these findings highlight that young males themselves have competitive spending motive and are motivated to earn higher incomes by exacting their efforts. These findings also suggest that the rising sex ratio in China may also have contributed to China's high GDP growth through competitive consumer spending and labor supply by young males.

Chapter 5. Conclusion of Thesis

Three essays in this thesis examine the connection between household consumption/finance and macroeconomic performance. The first essay finds adding housing consumption to the consumption bundle in a production economy has effect of lowering asset risk premia. In addition, it also finds that lower housing supply elasticity predicts lower equity risk premium and higher housing risk premium. The second essay shows that lower housing supply elasticity results in higher proportion of financial investment in stocks, and hence provides explanations for the geographic variations in household portfolio composition. Finally, the third essay finds that higher sex ratio gives rise to higher credit card spending and more worker efforts by young males. It suggests that the high GDP growth in China may be attributable to the competitive spending and labor supply behavior of young males induced by sex ratio imbalance.

These findings not only advance our understanding of the interactions between household behavior and macroeconomy, but also could provide practical implications for household decisions. For instance, the results from the first essay suggest that we should lower our expectation of equity risk premium if we expect that housing supply would become less elastic. For instance, in the case of China, we may expect that the aggregate housing supply elasticity would become lower in the future than today due to gradually exhausted developable land and population concentration driven by urbanization. Therefore, we may foresee a rise of aggregated stock price index due to lower aggregated housing supply elasticity and declining equity risk premium. The dependence of optimal proportion of investment in stocks on local housing supply elasticity shown in the second essay would suggest that households should adjust their investment strategies accordingly if they need to move to other locations. The second essay also provides evidence that there is a promising demand for instruments to hedge against housing consumption risk, an information useful to financial sector and government regulators.

Last but not the least, there are several limitations in the current research. For instance, the role of housing as collateral is absent in the theoretical models in the first two essays. Since the collateral effect has been found important in affecting asset risk premia in a market with incomplete risk sharing, extension of current models with heterogeneous agents and housing collateral to check the robustness of current model implications would be interesting. While the housing in the theoretical model of the second essay mainly serves as a consumption goods, an extension of the model by allowing housing investment (housing as an asset) is also interesting. In addition, although the results in the second essays show that households have demand for financial instruments to hedge against housing consumption risk, the minimal trading activity of the CME S&P/Case-Shiller HPI futures since the initiation of trading in 2006 suggests that improving the attractiveness of a financial product to households is always challenging and would be a promising area for future research.

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Appendix A: Log Approximation of the Stochastic Discount Factor

This appendix shows steps of deriving (10) based on (9) in Chapter 2. Taking the log of both sides of equation (9) in Chapter 2, we have

$$\ln\left(M_{t+1}\right)=\ln\left(\beta\right)-\gamma\ln\left(C_{t+1}\right)+\frac{\left(\sigma-\varepsilon\right)}{\sigma\left(\varepsilon-1\right)}\ln\left(\omega+(1-\omega)\left(\frac{C_{t+1}}{H_{t+1}}\right)^{\frac{1}{\varepsilon}-1}\right)+\gamma\ln\left(C_{t}\right)-\frac{\left(\sigma-\varepsilon\right)}{\sigma\left(\varepsilon-1\right)}\ln\left(\omega+(1-\omega)\left(\frac{C_{t}}{H_{t}}\right)^{\frac{1}{\varepsilon}-1}\right)$$

(a1)

Use the small case to denote the log of a variable such that $c_t-h_t=\ln\left(\frac{C_t}{H_t}\right)$, we can approximate $\ln\left(\omega+(1-\omega)\left(\frac{C_{t+1}}{H_{t+1}}\right)^{\frac{1}{\varepsilon}-1}\right)$ around c_t-h_t to obtain:

$$\ln\left(\omega+(1-\omega)\left(\frac{C_{t+1}}{H_{t+1}}\right)^{\frac{1}{\varepsilon}-1}\right)=\ln\left(\omega+(1-\omega)\left(\exp\left(\left(\frac{1}{\varepsilon}-1\right)\left(c_{t+1}-h_{t+1}\right)\right)\right)\right)$$

$$\approx\ln\left(\omega+(1-\omega)\left(\frac{C_{t}}{H_{t}}\right)^{\frac{1}{\varepsilon}-1}\right)+\left(\omega+(1-\omega)\left(\frac{C_{t}}{H_{t}}\right)^{\frac{1}{\varepsilon}-1}\right)^{-1}(1-\omega)\left(\frac{C_{t}}{H_{t}}\right)^{\frac{1}{\varepsilon}-1}\frac{1-\varepsilon}{\varepsilon}\left(c_{t+1}-h_{t+1}-\left(c_t-h_t\right)\right)$$

(a2)

Plug (a2) into (a1) and simplify, we can have:

$$m_{t+1}=\ln\left(\beta\right)-\gamma\left(c_{t+1}-c_t\right)-\frac{\left(\sigma-\varepsilon\right)}{\sigma\varepsilon}\left(1+\frac{\omega}{1-\omega}\left(\frac{C_t}{H_t}\right)^{1-\frac{1}{\varepsilon}}\right)^{-1}\left(c_{t+1}-h_{t+1}-\left(c_t-h_t\right)\right)\ \text{(a3)}$$

Because the intratemporal optimization implies a one-to-one mapping between the price of housing service (relative to numeraire good price which is normalized to one) and the relative quantity $D_t^h = \frac{\partial U\left(G_t\left(C_t, H_t\right)\right)}{\partial H_t} \Big/ \frac{\partial U\left(G_t\left(C_t, H_t\right)\right)}{\partial C_t} = \frac{1-\omega}{\omega}\left(\frac{H_t}{C_t}\right)^{-\frac{1}{\varepsilon}}$, we can replace $\frac{C_t}{H_t}$ in (a3) by $= \left(\frac{\omega}{1-\omega}\right)^{\varepsilon}\left(D_t^h\right)^{\varepsilon}$ to obtain equation (10) in Chapter 2.

Appendix B: Outline of Extended Model with Growth in Chapter Two

This appendix outlines an extended model with growth. The qualitative and quantitive properties of the model will be explored in future research.

The production technology in the non-housing sector is represented by the AK production function with labor-augmenting technology:

$$Y_t = (A_t)^{1-\alpha} \left(K_{t-1}^{b}\right)^{\alpha}, \quad 0 < \alpha < 1 \tag{b1}$$

The productivity level, instead of being stationary, is assumed to have stochastic growth:

$$\ln(A_t) = g + \ln(A_{t-1}) + \epsilon_t, \qquad \epsilon_t \sim N(0, \sigma_{\epsilon}^{2}) \tag{b2}$$

The technology process (b2) implies that the expected gross growth rate of A_t is $a_t = \exp(g + \frac{1}{2}\sigma_{\epsilon}^{2})$.

To ensure that the growth of housing consumption is the same as numeraire consumption and capital so that the balance growth path exists, the production of housing service is also subject to the aggregate productivity shock:

$$H_t = A_t^{1-\varphi} \left(K_{t-1}^{h}\right)^{\varphi} \tag{b3}$$

Given (b1) ~ (b3) and the nonseparable utility specification of (7) and (8) in Chapter 2, the agent's optimization problem in the growth economy can be summarized as:

$$\max_{\{C_t, K_t^b, K_t^h\}_{t=0}^{\infty}} E_0 \left(\sum_{t=0}^{\infty} \beta^t \frac{\left(\omega C_t^{1-1/\varepsilon} + (1-\omega) H_t^{1-1/\varepsilon} \right)^{\frac{\varepsilon(1-\gamma)}{\varepsilon-1}}}{1-\gamma} \right)$$

$$s.b.t.: \qquad (b4)$$

$$C_t + K_t^b + K_t^h = A_t^{1-\alpha} \left(K_{t-1}^b \right)^{\alpha} + (1-\delta)\left(K_{t-1}^b + K_{t-1}^h \right)$$

where H_t is represented by (b3).

Since all the variables in (b4) grow at the same rate, the model can be transformed by deflating these variables by A_t. By defining $C_t = \frac{C_t}{A_t}$, $H_t = \frac{H_t}{A_t}$, $K_t^b = \frac{K_t^b}{A_t}$, and $K_t^h = \frac{K_t^h}{A_t}$, the agent's optimization problem in the transformed stationary economy can be expressed as:

$$
\max_{\{C_t, K_t^b, K_t^h\}_{t=0}^{\infty}} E_0 \left(\sum_{t=0}^{\infty} \beta^t A_t^{1-\gamma} \frac{\left(\omega C_t^{1-1/\varepsilon} + (1-\omega) H_t^{1-1/\varepsilon} \right)^{\frac{\varepsilon(1-\gamma)}{\varepsilon-1}}}{1-\gamma} \right)
$$

$$
s.b.t.: \tag{b5}
$$

$$
C_t + K_t^b + K_t^h = a_t^{-\alpha} \left(K_{t-1}^b \right)^{\alpha} + (1-\delta) a_t^{-1} \left(K_{t-1}^b + K_{t-1}^h \right)
$$

Solving (b5) and transforming the stationary variable back to variables in the growth economy will allow us to examine whether connection between the elasticity of housing supply and asset risk premia still holds while matching import stylized facts about business and housing cycles.